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WORKPLAN AND METHODOLOGY

TASK I - DEVELOPMENT OF AN EQUITABLE MEDICARE ECONOMIC INDEX: AN ANALYSIS OF VARIATIONS IN PRACTICE COSTS AMONG SPECIALTIES

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WORK PLAN FOR TASK I ANALYSIS

I. Overview

During the late sixties and early seventies, physician fees for services to Medicare patients increased at a sufficiently rapid pace to warrant the attention of both Congress and the Administration. In 1972, this attention produced an amendment to the Social Security Act, designed to control the rate of growth in these expenditures for physician services. The proposed method of controlling fees focused on restricting the rate of increase of the prevailing charge rates, which are the maximum allowable reimbursement rates allowed physicians for Medicare services. Up to that time, increases in actual charges had determined the growth in the prevailing charge rates, but with this amendment, the growth in these rates was to be determined independently by a newly specified Medicare-Economic Index and not by the growth in actual charges.

The prevailing charge rates are a part of the "usual, customary and reasonable" (UCR) charge rate system which sets reimbursement rates for Medicare physician services. Under this system, the physician receives the lowest of the actual, customary or prevailing charges given certain reasonable allowances for variation in the service provided. The customary charge (Level I) is the physician's usual charge or more specifically the median charge for that service during the preceding screen year. Customary charges lag current charges by at most eighteen months due to the lag in

sampling for the screen year.

Prior to the introduction of the Medicare Economic Index, Medicare carriers calculated the prevailing charge (Level II) for a given service by setting it equal to the 75th percentile of all customary charges, each charge weighted by its frequency. The carrier grouped customary charges for similar specialists within a given locality. Under this system, increases in the actual charge increased the customary charge, albeit with a lag, and also increased the prevailing charge if most physician charges were growing as indeed has been the case. Thus, the system had a built in feedback link for inflation in actual charges. Prevailing charges, which had been designed as the usual ceiling for reimbursement, were reflecting steady increases in practice charges. The Medicare Economic Index was introduced in 1976 to bound these increases in prevailing charges by specifying an independent growth rate for charges based on cost and thereby breaking the connection between the changes in actual and prevailing charges.

In its form, the Medicare Economic Index resembles a forward Laspeyres price index in which the ratio of total price change (current to base period) equals the sum of the changes in the input price ratios (also current to base period), each ratio weighted by its own share of base period expenditure. If the input price ratios measure the change in the opportunity cost of factor inputs to the physician firm, then their properly weighted sum, which is the Index itself, represents the percentage change

in the cost of physician output. Each base-prevailing-charge-rate is then multiplied by the Index. If only one Index is calculated, each prevailing charge increases by the same percentage and the relative position of each charge remains the same. Thus far, due to the absence of detailed data, only a single national Index has been computed annually for updating the charges of all specialties in all regions. Also due to the absence of data, BLS subindices acted as proxies for the actual input prices with cost share weights coming from a separate survey. In this formulation and particularly with these data, the Index has decisively broken the feedback link from actual charges to prevailing charges.

The following discussion will review the organization for Task I analysis, whose work product will assist in the further development of the Index. The discussion outlines the proposed work in each of these areas:

1. the specification of the form of the Index;
2. the estimation of the weights for the Index;
3. the exploration of the relationships among cost components; and
4. the determination of separate specialty and regional groupings;

During the first project year, preliminary work will also be undertaken in the following areas:

- the selection of prices or price proxies, and the comparison of BLS indices and comparable survey data; and
- the examination of the impact of the Index on the physician's choice of inputs to the firm and services to Medicare patients.

II. Detailed Work Plan

1. Price Indices

An index number measures the change in the magnitude of a variable from one situation to the next. These situations usually consist of different dates but can consist of different places or different groups. The variable is usually but not always a composite of other variables. The most commonly referenced type of index is the price index and the most widely cited of the price indices is the BLS consumer price which is a forward Laspeyres index. As indicated above, the Medicare Economic Index has the form of a Laspeyres price index.

The forward Laspeyres price index is the ratio of the sum of current prices (p_{1i}), each weighted by base quantities (q_{0i}) to the sum of base prices (p_{0i}), each weighted by base quantities (q_{0i}). Stated otherwise, the Laspeyres index equals the sum of the ratio of current (p_{1i}) to base prices (p_{0i}), each ratio weighted by its own share of base period expenditure:

$$\frac{\Sigma\, p_{1i}\, q_{0i}}{\Sigma\, p_{0i}\, q_{0i}} = \Sigma\, \frac{p_{1i}}{p_{0i}}\, \frac{p_{0i}\, q_{0i}}{\Sigma p_{0i}\, q_{0i}} \quad \text{LASPEYRES}$$

A closely related alternative to the forward Laspeyres is the forward Paasche index in which the base quantity weights are replaced by current quantity weights:

$$\frac{\Sigma\, p_{1i}\, q_{1i}}{\Sigma\, p_{0i}\, q_{1i}} = 1 \Big/ \left(\Sigma\, \frac{p_{1i}}{p_{0i}}\, \frac{p_{1i}\, q_{1i}}{\Sigma p_{1i}\, q_{1i}} \right) \quad \text{PAASCHE}$$

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The Paasche index requires the collection of data for current period weights in each successive period; on the other hand, the Laspeyres index only requires base period weights. Consequently, constructing a Paasche index entails considerably more effort and the resulting survey costs usually limit its use. However, cost considerations address only one criterion by which to evaluate the relative merit of different index forms.

The Laspeyres and Paasche indices are only two of the infinitely large class of index number forms. To select a price index from members of this class requires a set of criteria to evaluate both their measurement accuracy and their correspondence to economic assumptions. Irving Fisher (1922) devised two categories of three tests each to assess the measurement accuracy of index numbers. These tests will be reviewed briefly and applied to the Laspeyres index, the Paasche index, and Fisher's ideal index, which is the geometric mean of the first two. The Fisher index has the form

$$\left(\left(\frac{\Sigma p_{1i} q_{0i}}{\Sigma p_{0i} q_{0i}} \right) \left(\frac{\Sigma p_{1i} q_{1i}}{\Sigma p_{0i} q_{1i}} \right) \right)^{\frac{1}{2}}$$

Fisher's first category included three simple tests of measurement consistency. The first of these is the identity test which requires that if a situation (e.g. a year) is compared to itself, it equals one. The second test is the units test which states that a change in the measurement unit of money will not change the value of the index itself. The third is the homogeneity or proportionality test which requires that if all prices change by an equal proportion from base to current period, the index itself will change by that

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proportion. (This test is quite useful for it is consistent with the property of demand functions that they are homogenous of degree zero in both prices and income.) The Laspeyres, Paasche, and Fisher-ideal index numbers all pass each of these three tests. The demonstrations are straightforward.

The second category of Fisher's tests determines whether an index number is transitive. The first of these tests requires that the forward price (quantity) index equals the inverse of the backward price (quantity) index, or in other words, that the product of the index measured from current to base and of the index measured from base to current equals one. This test insures that a change in the reference point of measurement does not change the order of the index number values. Satisfaction of this consistency test would be particularly important if the index were comparing groups; if the test were not satisfied, switching from one reference group to another might not preserve the ordering of the index number values. Among the index forms being considered, only the Fisher ideal index satisfies this test (the component subscripts having been dropped for clarity):

$$\left(\frac{\Sigma p_1 q_0}{\Sigma p_0 q_0} \quad \frac{\Sigma p_1 q_1}{\Sigma p_0 q_1} \right)^{\frac{1}{2}} \left(\frac{\Sigma p_0 q_1}{\Sigma p_1 q_1} \quad \frac{\Sigma p_0 q_0}{\Sigma p_1 q_0} \right)^{\frac{1}{2}} = 1$$

However, neither the Laspeyres nor the Paasche index numbers satisfy this test; in the forward Laspeyres case:

$$\left(\frac{\Sigma p_1 q_0}{\Sigma p_0 q_0} \right) \left(\frac{\Sigma p_0 q_1}{\Sigma p_1 q_1} \right) \gtrless 1$$

(i.e. not necessarily equal) because the inverse of the forward Laspeyres is

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the backward not the forward Paasche, the latter appearing as the second term above.

The second test in this category requires a transitivity across components. In the case of a price index measured across time, this test specifies that the index measured from the base to the current period equal the product of the same index measured from the base to an intermediate period and from the intermediate period to the current period. None of the indices being considered satisfies this test; in the Laspeyres case:

$$\left(\frac{\Sigma p_t \; q_0}{\Sigma p_0 \; q_0}\right) > \left(\frac{\Sigma p_k \; q_0}{\Sigma p_0 \; q_0}\right)\left(\frac{\Sigma p_t \; q_k}{\Sigma p_k \; q_k}\right)$$

(The inequality takes account of substitution due to price changes; the right hand side allows some substitution and the left none.) Of Fisher's six tests this is the most difficult one to satisfy. More significantly, the test is somewhat unrealistic in that it supposes an index could be developed which would use only current and base prices to measure the full price change over the entire period. An example might explain this objection to the test: if the values in a price vector shift from the base to the intermediate period (k) and then shift back to the base vector values in the current period (t), then the left hand side above (whether Laspeyres, Paasche or Fisher-ideal) would equal unity-would show no change-but would not necessarily equal the right hand side; the right hand side would reflect the effect of the sequence of interim price changes.

This test has a simple but important consequence for inter-group comparisons.

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In comparing group A to group C, one should not derive that comparison by the indirect route of comparing A to B and then B to C if the index number chosen fails this test. The test has another more important consequence and that is pointing to the desirability of taking into account the interim sequence of price changes from the base to the current period (e.g., calculating the right not the left hand side above). Doing this instantaneously would produce the only well-known index which satisfies this test: the continuous Divisia index. Its discrete usable counter-part is the chained Laspeyres index which has been shown above not to satisfy this transitivity test.

The final Fisher test requires that the product of the price and quantity indices equals the change in total expenditure. Neither the Laspeyres nor the Paasche indices satisfy this test: in the case of the Laspeyres:

$$\left(\frac{\Sigma p_1 q_0}{\Sigma p_0 q_0}\right)\left(\frac{\Sigma p_0 q_1}{\Sigma p_0 q_0}\right) \gtrless \frac{\Sigma p_1 q_1}{\Sigma p_0 q_0} .$$

However, the Fisher ideal index does

$$\left(\frac{\Sigma p_1 q_0}{\Sigma p_0 q_0}\cdot\frac{\Sigma p_1 q_1}{\Sigma p_0 q_1}\right)^{\frac{1}{2}}\left(\frac{\Sigma p_0 q_1}{\Sigma p_0 q_0}\cdot\frac{\Sigma p_1 q_1}{\Sigma p_1 q_0}\right)^{\frac{1}{2}} = \frac{\Sigma p_1 q_1}{\Sigma p_0 q_0}$$

This test is often referred to as the factor reversal test; note that the test requires that the total expenditure change is explained as the product of the price and quantity indices. This consistent separation of price and quantity changes can have useful applications if one is interested in isolating the price change component.

Other indices satisfy these tests. Stuvel (1957) has suggested a formulation using only the forward Laspeyres price and quantity indices. His derivation started with a separation of the change in total value ($\Sigma p_{i1}q_{i1} - \Sigma p_{i0}q_{i0}$) into price and quantity changes; in this derivation, Stuvel averaged prices in the quantity difference equation and averaged quantities in the price equation. This is the essence of the difference between the Fisher ideal index and the Stuvel index although he did not highlight it. Whereas Stuvel weighs each price difference ($p_1 - p_0$) by its own average quantity ($q_1 + q_0$)/2), Fisher's index weighs each price ratio (p_1/p_0) by the product of its base (current) expenditure share and the total Paasche (Laspeyres) price change. Thus, the Fisher ideal index weighs the individual price ratios by the product of own expenditure shares and the change in overall prices. The Stuvel index separates price and quantity changes explicitly (denoted I_p and I_Q respectively below):

$$I_Q = \frac{L_Q - L_P}{2} + \left(\left(\frac{L_Q - L_P}{2}\right)^2 + \frac{\Sigma p_{i1} q_{i1}}{\Sigma p_{i0} \Sigma q_{i0}}\right)^{\frac{1}{2}}$$

$$I_P = \frac{L_P - L_Q}{2} + \left(\left(\frac{L_P - L_Q}{2}\right)^2 + \frac{\Sigma p_{i1} q_{i1}}{\Sigma p_{i0} \Sigma q_{i0}}\right)^{\frac{1}{2}}$$

where L_Q is the Laspeyres quantity index and L_P the price index. I_p is the total effect of all price changes and I_Q is the total effect of all quantity changes. Stuvel showed that his indices satisfied all the tests which Fisher's ideal index satisfied; in particular note that $I_Q I_P = (\Sigma p_1 q_1)/(\Sigma p_0 q_0)$.

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The above analysis applies several measurement tests to selected price indices. The remainder of this Section evaluates the economic implications of some indices. The basic Task I objective is the development of a price index for physician firms. If a cost function exists for physician firms, then the price index should reflect the structure of the cost function. Unfortunately, only a small portion of the literature on price indices analyzes the correspondence between cost functions and price indices. Two exceptions are Vartia (1976) and Diewert (1976, 1978).

A basic observation about the Laspeyres index is in order before considering the work of Vartia and Diewert. In the Laspeyres index, the total price change equals the ratio of the sum of the products of base quantities and current prices to the sum of the products of base quantities and base prices. The use of base quantities as weights in the current period implies that no substitution occurs in production as a result of the price changes. This implication points to a very restrictive assumption about the nature of the underlying production function; namely, that at least in the short-run (between the base and current periods), the elasticity of substitution between inputs is zero because the production function is characterized by fixed factor coefficients.

On the other hand, comparing current prices and quantities to base quantities and current prices, the Paasche index asks the hypothetical questions: what would have been the values of production at base prices given the current mix of quantities? The Paasche index assumed that complete substitution among quantities has occurred. Consequently, the Fisher index, which is the product of the Laspeyres and Paasche, brackets the substitution possibilities. If no substitution occurs, then the Fisher index reduces to the Laspeyres and with complete substitution to the Paasche.

Diewert developed the link between cost functions and the Vartia and Törnquist indices. Diewert's development is rather technical and only some results will be described. He defined a general criterion for the equivalence of the price index and the ratio of the cost function evaluated at current prices to the cost function evaluated at base prices. This criterion stated that the cost function is exact for a price index if given the base and current price vectors, the dual linearly homogeneous aggregator (i.e., 'production') function has a maximum which corresponds to the input base and current quantity vectors respectively.

Diewert then demonstrated that the Vartia index is exact for the Cobb-Douglas unit cost function. The Vartia index is defined as

$$\ln P_v = \Sigma \frac{(p_{1i}q_{1i} - p_{0i}q_{0i}) / (\ln(p_{1i}q_{1i}) - \ln(p_{0i}q_{0i}))}{(\Sigma p_{1i}q_{1i} - \Sigma p_{0i}q_{0i}) / (\ln(\Sigma p_{1i}q_{1i}) - \ln(\Sigma p_{0i}q_{0i}))} \ln \frac{p_{1i}}{p_{0i}}$$

and the Cobb Douglas unit cost function is defined as

$$\ln c = a_0 + \Sigma a_i \ln p_k$$

where c is cost per unit, a_0 a constant and p_i the input prices and $\ln P_v = \ln c_1 / \ln c_2$. The Vartia quantity index only substitutes the ratio of quantities in the cost term. The product of the Vartia price and quantity indices equals the change in total cost; in other words, the Vartia indices satisfy the Fisher factor reversal test. Furthermore, if the firm's technology corresponds to a linearly homogenous Cobb-Douglas production function (i.e., the elasticity of substitution equal to one), then the use of the Vartia index insures that the price ratios are weighted by factors consistent with the underlying tech-

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nology. With Cobb-Douglas technology, use of the Laspeyres index would over-report the price change while use of the Paasche index would under-report it.

Diewert goes on to show that the Törnquist index is exact for the translog unit cost function. This result is quite interesting because the translog function is much more flexible than the Cobb-Douglas, allowing non-constant elasticities of substitution. (As Theil (1973) noted the index is not unique to Törnquist and is frequently used as the discrete approximation to the Divisia index.) The Törnquist price index is

$$\Sigma \tfrac{1}{2}\left((p_{i0}\, q_{i0}/\Sigma\, p_{i0}q_{i0}) + (p_{i1}\, q_{i1}/\Sigma\, p_{i1}\, q_{i1}) \right) \ln\, (p_{i1}/p_{10}).$$

The first two ratios are the expenditure shares in the base and current period respectively. The translog unit cost function corresponding to a linearly homogeneous translog production function is

$$\ln c = b_0 + \Sigma\, b_i \ln p_i + \tfrac{1}{2} \Sigma\Sigma b_{ij} \ln p_i\, p_j$$

where c is cost per unit and p_i the input prices. The price index equals the ratio of the unit cost function evaluated at current prices to it evaluated at base prices or if this price index is designated as $\ln P_T$, $\ln P_T = \ln C_1/\text{Ln } C_0$, where C_1 is evaluated at current prices and C_0 as base prices. The quantity index substitutes quantities for prices in the price ratio terms above but the expentiture shares are identical. The quantity index equals the ratio of the aggregator function evaluated at current quantities to the aggregator function evaluated at base prices.

The product of the price and quantity indices does not quite equal the change in total cost but as Theil (1973) showed, approximate it to the third order of smallness. This means that the Törnquist index does not quite satisfy the Fisher factor-reversal test. If the firm's technology corresponds to a linearly homogenous translog production function, then use of this index will insure that price ratios are weighted by factors consistant with the underlying technology. The translog function allows a very flexible specification for the technology, and includes the Cobb-Douglas as a special case (σ=1).

Diewert demonstrated the exact correspondence between the Törnquist index and the translog function under constant returns to scale. Diewert (1976) also showed that when the translog does not exhibit constant returns, the Törnquist index does not correspond uniquely to the translog but to a wider class of functions including the translog. In this general form of the translog, output enters not as y_0 in the base and y_1 in the current period, but as $(y_0 y_1)^{\frac{1}{2}}$ in both. (Diewert demonstrated this result for the more general non-homothetic case in which the "cross-effect" of input price with output enters the cost function.)

Note that the index has two other rather minor limitations. First, if a quantity is present in some periods but not others, the log operator in the quantity index will be undefined in years in which this quantity is absent. This problem does not exist in the price index. Secondly, unlike the Laspeyres, Paasche, and Vartia, this index is not quite consistent in aggregation: combining subindicies will not exactly yield the overall index.

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The index has a major non-technical limitation. It requires the estimation of base and current expenditure shares because the weights are averages of the two periods. Continual sampling to determine current expenditure shares entails considerable re-survey expense; as indicated above, the Paasche index encounters the same objection. Of course, use of the information from the last and current period avoids the "drifting" problem which occurs when only a Laspeyres index is used; the standard resolution for the problem is the chaining (i.e., multiplying) of up-dated Laspeyres indices together. Chaining the Törnquist indices would improve this solution.

Diewert's work suggests a way to avoid some of the costs of resurvey. If the underlying technology remains constant then the estimated coefficients of the translog unit cost function can predict the Törnquist index. The predicted Törnquist index would equal the ratio of the product of the estimated coefficients $\hat{\beta}$ and the current price vector P_1 over the products of the coefficients and the base price vector P_0. For this reason, we will estimate cost function parameters; the discussion concerning that estimation is in the next section. Without the same re-sampling which this procedure seeks to avoid, one could not be confident that the underlying technology had not changed and that the estimated coefficients used in the ratio's numerator were correct. However, one might be somewhat confident that a substantial shift would not occur within, for example, five years which is the usual period for changing the base in the use of the Laspeyres indices. In fact, use of the annual re-survey data from the physician's practice cost survey to estimate cost functions for sequential years may suggest whether significant changes in the elasticities of substitution will have occurred in the near term. Review of the literature may also assist this evaluation.

However, if the possibilities for substitution among inputs in physician firms is not considerable, then the correspondence between the flexible translog and the Törnquist index may not be that important. Through our own estimation and particularly that of others we will attempt to determine some range on the elasticities of substitution. Note that if such substitution was very low or very slow, then use of a Laspeyres type index would not mis-respresent critical cost shares. As part of this work, we will attempt to compare the implicit substitution possibilities represented by the Törnquist index with that represented by the Fisher ideal and the Stuvel indices by using some of Uzawa's (1962) results representing the Allen partial elasticities of substitution in terms of cost function derivatives.

Up to this point, the discussion has reviewed both measurement and economic criteria by which to evaluate various index forms. These two criteria address the issue of consistency in indices. Furthermore, the economic discussion focused on the extent to which the price index weights accurately represented the factor substitution possibilities. This discussion reaches the issue of efficiency but not that of equity and the Task I subject is the formulation of an equitable Medicare Economic Index. To reach this issue requires some consideration of the meaning of equity.

In a practical context, which avoids abstract arguments about welfare functions, equity might merely mean the absence of discrimination amoung groups differing significantly in technology. In this case, the groups might well be the separate specialties which would be treated alike only if their production

technology, as measured by their estimated elasticities of substitution, were not significantly different. If for some a priori reason particular groups were known to be inefficient in the use of inputs so that the estimated coefficients did not reflect an efficient technology, then consideration could be given to the use of other proxies for the true parameters in order ultimately to encourage efficiency. The application of this equity rule is then consistent with the efficient use of inputs in production. To an economist, this application would seem desirable because it represents the movement toward or preservation of Pareto optimality (recognizing of course the considerable market imperfections in the area of physician firms which point to second-best solutions).

On the other hand, establishing equity might mean not only avoiding discrimination among groups but also limiting the growth rate in Medicare fees to the growth rate of prices in another sector. This seems to be the equity goal because of the method chosen to define the fee limits through the Index. The current formulation of the Medicare Economic Index uses the growth rate in input prices from other sectors as a limit for the implied growth rate in our sector input prices. Through the Index, this limit on input prices translates into a limit on fees. How effectively the chosen means will reach the implicit goal is not clear and we will examine this issue in general terms. This examination will consider possible goals behind the index: for example, the maintenance of parity in the growth rate between total Medicare physician expenditures and earned income. Finally, the goal of the index may not be explicit parity between growth rates but may merely be the establishment of some limits by which to control growth in fees and this possibility will be included in the discussion.

2. Estimation of the Weights

The previous Section reviewed the structure and the merits of several indices. Every index requires a set of prices and estimated cost shares, and this Section addressed the problem of the estimation of the cost shares.

With the Laspeyres price index, only an estimate of the base expenditure or cost shares is necessary because the Laspeyres assumes the unimportance of the substitution between the base and current periods. On the other hand, the Paasche index needs the estimate of current cost shares. The Fisher Ideal, Stuvel (implicitly), Vartia and Törnquist indices all require estimates of cost shares in current and base periods. As indicated at the end of the last Section, estimation of the Cobb-Douglas and translog unit cost functions provide indirect measures of the Vartia and Törnquist indices respectively. The discussion in this Section includes a description of the proposed cost function estimation.

One can approach the estimation of cost shares directly by calculating the share of total expenditures in each of the several categories of input cost. The Physician Practice Cost Surveys in 1975 and 1976 include physician firms' annual expenses for all major input categories: office and auto expenses, medical equipment and materials, malpractice insurance, non-physician labor and physician net income. (The 1975 Survey contains much more detailed administrative expenses; and the Abt report has extensively analyzed these expenses.)

For all specialties, the tables in Appendix I present the cost shares of these categories excluding physician net income. Each ratio entry in the tables is

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the dollar amount expended in the indicated category over the total dollar amount expended in all categories, excluding non-physician income. Consequently, each ratio is the sum of costs in the indicated category to the sum of all non-physician costs where each sum is taken over all physicians in a given specialty. These ratios weight the cost distributions of the larger firms more heavily and therefore have bias positively related to the firm's total cost. A more appropriate ratio estimator is the average of the ratios, rathern than the ratio of the averages, because the former does not weigh firms differently. If within a given specialty, the population (not the sample) cost share parameters differs due to size of output, different prices, etc., then only regression analysis can capture these differences and that approach is introduced below. Note that these ratios presented in the Appendix tables and the chi-square statistic were intended to be only convenient and descriptive of the differences among specialties.

The appropriate test to determine simultaneous equality of the cost share patterns across specialties is a multivariate analysis of variance. (Anderson, 1958) This test consists of multiple analyses of variance, each analysis of variance testing the equality of a given mean (e.g., a particular type of cost share - labor, materials) across all specialties and the multiple analysis testing the equality of all these sets of means simultaneously. (Because the n cost shares sum to one, testing the equality of the n-1 shares is sufficient.) Therefore, using a multiple analysis of variance would test whether significant differences existed in the distribution of all cost ratios across all specialties. Performing the test directly on the ratios is desirable because the ratios enter directly into the price index formulation. However, ratios can be biased estimators of the relation between the two variables if a regression line between the two variables does

not pass through the origin but has a constant term. This possibility might well occur in the relation between a cost component and the total cost; when the total is positive, some components can indeed be zero (if those inputs are unused in production).

The problems with ratio estimators leads to a consideration of alternatives. The next most naive approach is to average the cost components per physician across all firms in a given specialty and test whether the distribution of these average costs differed across specialties, again using a multivariate analysis of variance. If opportunities for economies of scale were constant across firms and related to the number of physicians per firm yet this number on average differed across specialties, then the specialities with larger firms would show smaller cost component averages, ceteris paribus, but the ratios would be identical across firm size because the distribution of cost across categories has by assumption not changed. However, even this test on the component averages would not locate the cause of inter-firm or inter-specialty differences; in the example given, all-around lower costs could have resulted from lower prices. Establishing a causal relationship between changes in total costs (or its components) and other factors requires introducing a model to explain variation in form costs.

An appropriate model is the cost function of the firm which relates total cost to output and the set of input prices. The cost function contains the same information about production as its dual the transformation function which describes the efficient production possibility set, if the cost function is non-decreasing and strictly positive when output is nonzero and if the cost function is concave and linearly homogeneous in it input prices for any given level of output.

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Several problems arise in the estimation of cost functions for physicians. If the firm does not cost minimize, then the estimated parameters of any given cost function will not represent the possibilities for substitution under conditions of efficient production. This outcome makes interpretation of the estimated parameters difficult if not impossible. Note that if the producer in a non-competitive industry is a cost-minimizer, identifying the parameters of the firm in the non-competitive case may still be a problem because equilibrium price and quantity will be a function of both supply and demand. Another way to view this problem is that the system is not identified by estimation of the cost function alone. A further problem is the endogeneity of the price of the physician's own labor; to interpret the estimated parameters as determining elasticities of substitution requires that the prices be exogenous to the firm. Similarly, output should be exogenous. In the case of physician firms, neither output nor price of physician labor is likely to be exogenous; in fact, they are probably jointly determined.

As a mathematical function, a cost function can assume a variety of forms. Two commonly used forms are the Cobb-Douglas and the translog. Each corresponds to different assumptions about the possibilites for substitution among factor inputs although their differences extend beyond this because the translog allows non-homothecticity and a greater flexibility in the form of non-constant returns to scale. The Cobb-Douglas form implies that the Allen partial elasticities of substitution between pairs of inputs ($_{ij}$) is constant and equal to one, while the translog form imposes no restrictions: the elasticities can assume any

real value and can vary with the level of the cost shares and the relationship between a cost share and other input prices. The greater flexibility of the translog usually recommends its use although its estimation is more costly. The following discussion will review both the forms of the Cobb-Douglas and the translog function, discuss some of the estimation problems and conclude with a few comments on the use of the cost functions in the price index.

The Cobb-Douglas production function has the form:

$$\ln Y = k + \Sigma \alpha_i \ln X_i$$

where Y is output, and the X_i the set of inputs. Cost minimization subject to this production function (and through a series of substitutions) yields a total cost function, whose estimated form is

$$\ln (C-P_j) = \alpha + \Sigma \alpha_i \alpha_Y \ln(P_i-P_j) + \alpha_Y \ln Y$$

where C is total cost, each P_i is the price of the input, each α_i is from the production or function and α_Y is the inverse of the sum of the α_i. If the sum of the α_i is less (greater) than one, decreasing (increasing) returns to scale are present. The product $\alpha_i\alpha_Y$ equals the cost share of the ith input (p_iX_i/C). In order that the estimated parameters of all cost shares sum to one $\Sigma p_iX_i/C = 1$, the estimation of the cost function must incorporate the equivalent restriction $\alpha_i\alpha_Y = 1$ which is implemented by subtraction of any one price from total cost and all other prices.

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Information on one or more prices may not be available; for example, the Survey does not have price data on office materials, equipment or floorspace or medical materials or equipment. If these prices were correctly assumed to be constant across firms, then their effect would enter the estimated constant term but since several prices are missing, their estimated cost shares would only be known in the aggregate as $1 - \Sigma \alpha_i \alpha_\gamma$ where $\Sigma \alpha_i \alpha_\gamma$ covers only those inputs with non-missing price data. In fact, the only prices available are the wage rates for labor. Thus, even if the unavailable prices were constant across firms, no unique cost shares could be inferred. (The Cobb-Douglas function has a redeeming feature here; its assumption of strong separability in inputs (all $\sigma_{ij} = 1$) is equivalent to the existence of a consistent aggregate subindex of prices so that if one price could be found which would describe the entire set of missing prices no information would be lost, but unforunately no such price index is in the Survey or conveniently recomends itself.) Because the missing prices are non-labor, one might be tempted to treat them all as "capital" and use the price of external financing as a proxy but this use confuses the real and the financial sides of the firm's operation; for example, if two firms faced the same price vector of inputs but different prices for external financing because of imperfect capital markets, then their choice of input might well differ but not because of different possibilities for substitution in production.

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The translog cost function has the form

$$\ln C = a_0 + a_Y + \tfrac{1}{2} g_{YY} (\ln Y)^2 + \sum_i a_i \ln P_i +$$

$$\tfrac{1}{2} \sum_i \sum_j g_{ij} \ln P_i P_j + \sum_i \sum_j g_{Yi} \ln Y \ln P_i$$

where C is total cost, Y output and P_i the input prices. For the translog function to have the desirable characteristic of homogeneity of degree one in prices, which implies that a uniform increase in all prices results in a proportionate increase in total costs, the following restrictions must hold:

$$\sum_i a_i = 1 \; ; \; \sum_i g_{Yi} = 0; \; \sum_i g_{ij} = \sum_j g_{ij} = \sum_i \sum_j g_{ij} = 0$$

These restrictions constrain the cost share sum to equal one (Christiansen and Greene, 1976).

Differentiating the cost function with respect to prices yields a set of cost share equations of the form

$$S_i = a_i + g_{Yi} Y + \sum_j g_{ij} \ln P_j$$

where the S_i are the cost shares of the i th input. Because of the probable substantial multicollinearity among all the price and cross-price variables in the cost function, estimation of the cost share equations yields considerably more precise parameters. However, the cost share equations do not include several parameters appearing in the cost function itself. Their absence prevents testing hypotheses of homotheticity and homogeneity which respectively imply $g_{Yi} = 0$ and $g_{Yi} = g_{YY} = 0$. If these hold and if all the $g_{ij} = 0$, then the translog reduces to the Cobb-Douglas form. Joint estimation of the cost function and the cost share equations allows the testing of these hypothe-

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ses. Because the cost shares must sum to one, one cost share equation must be dropped from this system. To derive parameter estimates which are invariant to the equation being dropped requires maximum-likelihood estimation; iterating the Zellner procedure (IZEF) is one approach (Christensen and Greene, 1976).

The cost function considered above included only one output. Multiple outputs may more accurately describe the physician firm. Specifying multiple outputs for the translog function requires adapting the log-operator of output because of the possibility that one or more of the outputs for some firms might be zero. Caves, Christensen and Trathaway (1978) have proposed use of the Box-Cox metric $f(y) = (y^k - 1)/k$ to replace the log metric. The log metric remains for prices and total cost allowing easier parameter restrictions to test homogeneity of degree one in prices. Their generalized translog multiproduct cost function (GTMCF) has the form

$$\ln C = a_0 + \sum_i a_i (Y_i^k - 1)/k + \sum_i b_i \ln P_i + \tfrac{1}{2} \sum_i \sum_j g_{ij} \ln P_i \ln P_j +$$

$$\tfrac{1}{2} \sum_i \sum_j d_{ij} (Y_i^k - 1)/k \, (Y_j^k - 1)/k + \sum_i \sum_j r_{ij} (Y^k - 1)/k \ln P_j$$

where the Y_i are outputs, P_j prices and C total cost. Non-linear estimation is required to determine which value of k minimizes the sum of squared errors.

Extensive data unavailability affects both the translog and the Cobb-Douglas but may limit the use of the computationally more expensive translog (and particularly the GTMCF) in the estimation of the cost function. Only prices for non-physician labor appear in the Survey. Estimation of a cost function with only the available prices will create serious econometric problems with omitted

variables; only if the omitted variables are uncorrelated with the included variables will the parameter estimates of the included variables not be biased. Assuming this lack of correlation, the estimated elasticities of substitution among labor inputs would be unbiased.

The estimate of the intercept will still be biased even if the omitted and included variables are uncorrelated, a particular problem in the translog cost share equation where the intercept identifies the impact of own price on the share; intercept bias will not disappear unless the omitted variable were irrelevant which by assumption it isn't (Kmenta, 1971). Only if the omitted prices were constant across all firms, which is extremely unlikely will their absence make no difference at all.

Some of the variables included in the Survey present problems. The first problem occurs with the specification of output. We will review the literature to assess which unitary measures of output have been used. The actual output is a potential change in the patient's health status, which is unmeasureable; the most commonly used output measure is simply the number of visits. Some investigators have used output measures such as total revenue which would be inappropriate in the cost function. In the estimation of the cost function, we will attempt to correct for some of the endogeneity of the output measure through two-stage least squares.

Disaggregating output could point to the construction of a case-mix index which recognized that each physician produces a different mix of diagnoses and treat-

ments. The supplementary work in Task 1 on Case-Mix Analysis indicates the direction of our exploratory efforts. This Analysis points to difficulties in linking the NAMCS data to the Survey data (especially via Survey question 35). Using a unitary measure of output which obscures the true variation in the multiple outputs may bias the estimated coefficients on the price of physician own labor because that price is probably positively correlated with case-mix components requiring more effort (e.g., more difficult diagnoses and treatments). The Case-Mix Analysis does not address one basic characteristic of both physician input and output; namely, its stochastic nature. Physicians can't accurately control the mix of problems presenting themselves or the outcomes from the treatments they recommend. The variation across specialties in malpractice expenses indicates among other things that uncertainty in treatment outcomes varies substantially. How this uncertainty affects the physician's production function is not clear.

Disaggregating output presents an interesting possibility for assessing the impact of the Index. The case-mix approach characterizes output as the mix of diagnoses and treatments. Another approach would be to characterize output as the mix of patient types, specifically Medicare and non-Medicare. If the Index creates an effective ceiling, then the coefficient on the cross-product term could indicate the extent to which substitution away from Medicare patients occurred.

The other major problem is the measurement of the price of physician labor. The price of physician labor is not exogenous to the firm; the physician who organizes a practice will decide simultaneously an expected output and an

al unit but measuring it is difficult. A proxy might be the fee for the most flexible of the services offered, the return visit, adjusted by its average length and reduced to exclude administrative expenses. In the long run, the private opportunity cost might be defined differently: the potential physician can choose among occupations and the practicing physician can choose between practice settings (e.g., a salaried government or HMO position would be alternatives).

Consideration of the correct price of physician time raises a related issue. That issue is the proper model for the physician firm. Some investigators have modeled the firm as if it were only an individual physician supplying own labor and others as if it were a firm. If the firm is the appropriate model, then estimation of the "supply of labor" function is a mis-specification even in the case of the solo-practitioner. The correct specification would be the firm's derived demand for physician labor function which should include firm output, the price of own input and the price of all alternative inputs for all observations even though some of the firms do not use some of the inputs. Determining which is the appropriate model is difficult; use of specification tests in estimation might help.

Another serious problem in the Survey is the extent of measurement error in the variables. The tables in Appendix I suggest the extent of the problem. In a regression, errors in variables create bias in the estimated coefficients unless in the limit the value of the term on the left vanishes:

$$\text{plim}(1/n(X'(u-V\beta))) = \text{plim}(1/n\ X'u) - \text{plim}(1/n\ \tilde{X}'V)\beta - \text{plim}(1/n\ V'V)\beta$$

where n is the sample size, X the measured and $\tilde{X}$ the true matrix of the values

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of the variables, u the vector of the residuals in the regression without error and V, the vector of the measurement errors (Johnston, 1972). Even if the first term on the left vanishes because u is uncorrelated with the true value of X and the measurement error ($X = \tilde{X} + V$), the third term will not vanish because the measurement error itself will be positive in the limit. Also, the second term will possibly be positive in the limit because the measurement error will likely be correlated with the true values of X; for instance, the measurement error in physician gross and net income probably increases with the true value of these variables.

The estimation of the translog or Cobb-Douglas cost function would be very useful because of their direct role as analogues to the Törnquist or Vartia price indices. Unfortunately, estimation of the appropriately specified cost function seems unlikely given the available data. Some alternatives might be investigated. The cost shares might be related to available measures of output, input prices, and other exogenous variables characterizing the structure of the local market for health 'production'. Such a set of equations would not correspond to a cost function but would specify a set of 'behavioral' cost share functions, wnose stability is unknown because of their essentially ad hoc character.

However, if variables are selected for their ability to explain variation in the cost shares, then those which enter will minimize the residual given the current characteristics of the physician, the practice and its locale. If the cost share estimates are derived as the product of the estimated coefficients and the characteristics of the individual practice, then these

estimates will reflect the existing local market. Thus, the more ad hoc the specification, the greater the likelihood that the cost share estimates will allow the Index to maintain market characteristics, including inefficiencies. For instance, if more non-competitive areas (e.g., those with both greater than average fees and physician density) showed cost shares weighted more heavily toward more costly inputs (e.g., physician input), then the Index would permit these areas a greater increase in fees. To remedy this problem would require dropping the variable from the estimation or substituting another value for the independent variable for these firms.

To draw the threads of this discussion to a conclusion, we propose a three-tiered method of determining the cost share weights. First, we will calculate the overall cost shares directly as averages of the individual shares, testing to determine if significant differences exist among specialties (c.f. Section 3). Second, we will estimate the Cobb-Douglas cost function with available prices constraining the labor shares to equal one less the shares of the unincluded variables and acknowledging the serious specification errors which omission of the other variables would create. We will also estimate the translog for limited portions of the sample. Third, we will estimate ad hoc cost share equations which include variables other than price and output.

Collection of the missing price data in future years would allow estimation and use of the cost function in the Index. However, determining the price of many inputs is difficult. An alternative would be to sample not the prices but the quantity of inputs and estimate the production function instead of the cost function. The price index could then be approximated indirectly

by using the result that the product of the price and quantity indices equals the ratio of the current to base expenditure or explicitly by calculating the ratio of two other ratios: the ratio of current to base expenditures over the ratio of current to base quantity index. For those indices which don't satisfy the factor reversal test exaclty (e.g., the Törnquist), this calculation would only be approximate.

3. Cost Component Relationships

The previous Section presented methods to calculate the weights used in the Medicare Economic Index. This Section briefly reviews related work aimed at further examination of the nature of the cost components. This work includes the examination of the levels of the components, their relation to net physician income, the rates of change in those levels and the relationship between other factors and both the levels and the cost shares.

For each specialty grouping, we will calculate the mean for each cost component per physician. The cost components will include the same categories as listed in Section 2: office and auto expenses, medical equipment and materials, non-physician labor, malpractice insurance, and physician net income. (These are the listed categories on the 1975 and 1976 Physician Practice Cost Survey.) Dividing each component by the number of physicians in the practice before averaging across all physicians in the same specialty will partially normalize the cost components by producing values which are costs per physician.

This normalization does not remove the effect of possible economies of scale. If opportunities for economies of scale were constant across specialties and related to the number of physicians in the practice but at the same time the average number of physicians per practice differed across specialties, then the costs per physician would also differ across specialties. However, if the analysis in Section 2 indicates what other researcher have shown, namely, that at most only modest economies of scale are present and are exhausted with fairly small practices, then one might apply the same analysis suggested in

the beginning of Section 2 to determine whether significant differences among specialties exist.

To determine the equality of each set of cost components simultaneously, we will again use a multivariate analysis of variance test. This approach will test a joint hypotnesis, a composite of several individual hypotheses. Each individual hypotheses specifies that the means for a given cost component are equal across all specialties. The test for that equality is a one-way analysis of variance. The test for the simultaneous equality of the means across specialties for the several cost components requires a multivariate analysis of variance. We will calculate the appropriate F-statistic to test whether significant differences exist (Anderson, 1958).

For all specialties grouped together, we will also perform a multivariate analysis of variance to determine if cost components vary across size of practice. We will measure size of practice by number of physicians and cost components by cost per physician. Again we will test a joint hypothesis, each individual hypothesis specifying that the means for a given cost component are equal across all practice sizes. This analysis will supplement the work in Section 2 and will help determine whether costs per physician decline in some intermediate range of practice size.

To examine the relationship of physician net income to total other costs, we could rely on a simple correlation between the two. We could posit a simple behavioral relation which would argue that if physician net income were under the firm's control, then it might be explained in a multiple regression by

output measures and total non-physician costs. We also can determine whether the relation between physician net income and total non-physician cost were correlated by using the Wilcoxon-rank test, the rank test being useful because of the possible bias in reporting errors for both costs. However, none of these ad hoc tests reaches the issue of whether high cost firms exist, firms whose costs are not justified by the level and mix of their output. A better approach would be to examine the pattern of the estimated residuals in the estimated total cost equation to determine if the size or variance of the estimated residuals increases for those firms having large total costs. If the translog functions were estimated, one could also examine the estimated residuals for these same firms to determine the pattern of estimated cost shares.

Both the 1975 and the 1976 Surveys contain questions relating to the physician firm's cost components. The question numbers and the cost categories to which they refer are the following:

	1975	1976
OFFICE EXPENSES	57-1,3,60,61	42-1
AUTO EXPENSES	57-2	45
MEDICAL EQUIPMENT	57-4	42-2
MEDICAL MATERIALS		42-3
MALPRACTICE INS.	62	44D
PHYSICIAN EMPLOYEES	68	48A
NONPHYSICIAN EMPLOYEES	68	53
TOTAL DEDUCTIONS	74	57,61
NET INCOME	62	62
GROSS INCOME	72,73	55,58

To measure the change in cost components for each specialty, we will compute both the change in individual cost components (75 to 76) divided by base year cost component (75) and this same change divided by base year total cost. The sum of the second percentage changes will equal the overall change in costs for a specialty. We can again apply a multivariate analysis of variance to the second percentage change. (Because the variable is a ratio, some caution should be exercised in interpreting the results.) To determine whether changes in cost components have been similar across components for any given specialty, we can we can apply a one-way analysis of variance to the set of changes for each specialty. A more complete test for structural shift in cost shares from 1975 to 1976 would utilize the Chow-test applied to the pooled estimation and to the separate estimation of cost functions for each year. We will consider the time series problems involved in pooling.

Finally, we will examine the impact of other factors on the estimation of cost shares. To do this, we will include other variables in the estimation of the translog cost share equations discussed in Section 2. These variables would include measures of the physician's individual background (e.g., age), the physician's practice (e.g., mixture of patient types, diagnostic services requested), or the area's characteristics in which the practice is located (e.g., measures of physician density by type of physician, number of hospital beds). In Section 2, we indicated that both output and the physician wage are endogenous to the firm, causing biased estimates of their coefficients in the cost share regressions. A potential remedy replaces the variables with their predicted values using two-stage least squares where excluded exogenous variables and included exogenous variables are used in the first stage to estimate the

predicted values. The excluded exogenous variables will include some of the variable indicated above. Entering these variables through the two-stage least squares process does not require making a priori assumptions about the strong separability of these variables and the other included variables in the cost equation itself. If the strong separability assumption is made, some of these variables could be separately included in the cost equation itself.

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4. Specialty and Regional Groupings

Currently the Medicare Economic Index does not distinguish among specialties or regions but consists instead of a single index with one set of cost shares. Using this single index and implicitly ignoring these differences among specialties creates potential problems because the Index may generate increases in prevailing fees which do not correspond to increases in cost, particularly for those specialties whose cost shares are far from the average.

However, if these inter- or even intra-specialty differences in cost shares do no reflect differences in technology, the estimation and use of separate cost share weights will reflect other influences. Estimating the parameters of the cost function from the survey data presents some considerable difficulties as indicated above in Section 2. The Survey data is deficient but even were the data complete and available, the resulting estimated parameters of the cost functions would represent the impact of market distortions. Work to date on testing the relative economic efficiency of firms has dealt with the case of allegedly undistorted markets in which all firms are price-takers (Yotopoulos and Lau).

In the apparent absence of a definable test to determine relative economic efficiency, we propose two ad hoc procedures. First, we will perform an analysis of covariance to test for significant differences among specialties. The analysis of covariance requires finding the estimated residual for the pooled sample (SSE_p) and for the classes within the sample ($\sum_{i=1}^{P} SSE_i$) and computing the following test distributed as an F-statistic with indicated degrees of freedom.

$$\frac{SSE_p - \sum_{i=1}^{p} SSE_i/k(p-1)}{\sum_{i=1}^{p} SSE_i/p(m-k)}$$

where k is the number of variables, p the number of classes, and $\hat{m}$ the number of observations per class. (The degrees of freedom specification is different if there are different numbers of observations per class.) This analysis of covariance will be applied directly to the estimated cost functions using the specialties as the classes. We will first test for overall homogeneity across all specialties. If that test indicates no significant differences across classes, then the testing would stop and it could be assumed that the single index is adequate. If significant differences are indicated, we will then divide the sample into surgical and non-surgical specialties, repeating the test for each of these two samples. If significant differences again exist, we will then divide the non-surgical sample into general/family practice and remaining specialties. The division into classes at each step should follow some notion of different practice organizations; for example, several surgical specialities may show no significant differences in cost shares and may be arguably treated as a single class. (Because of the sequential testing we will consider some adjustment to the degrees of freedom.)

Second, this test could be used within specialties if the specialties were partitioned into classes on the basis of cost per visit. If we accept the hypothesis that specialties with lower cost per visit constituted more "efficient" firms in some sense, then we could test for the differences in

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cost functions between the lower and higher cost firms. If no difference existed then the cost shares of the specialty (or group of specialties) could be used but if significant differences were indicated we could consider using the cost shares of the lower cost partition. These differences might be attributable to greater efficiencies in the lower cost firms or to other influences including market distortions.

Because of limited samples within several of the specialties, we will be unable to test for simultaneous homogeneity across region-specialty pairs unless we restrict those pairings to a reasonable minimum. A more inclusive set of pairings would be possible were we to pool the 1975 and 1976 Surveys which we will consider doing. On the other hand, we could allow the regional variations to enter only through price variation, which seems to be more plausable in the sense that within a given specialty or group of specialties the technology of physician practices should not vary substantially across regions. If we puruse this course, we may want to test to determine whether the sample shows significant differences among the regions in cost per visit functions, rather than total cost functions, by including regional dummy variables in this type of cost function. If significant differences appeared, that would warrant use of regional prices or use of adjusted national prices. One source for that adjustment might be the estimated coefficients themselves.

DATA ANALYSIS FOR PROJECT AREA I

Cost Shares Estimates

One objective of this project is the determination of the actual cost shares of each expense category facing the physician firm. These cost shares will function as the weights in the Medicare Economic Index.

The most fundamental approach to determine these cost shares is to calculate the ratio of each expense item to total expenses for each specialty. This method assumes that no significant variation in cost shares occur within the specialty and that significant variation does exist among specialties. To test the first assumption requires estimation of cost share equations as derived, for instance, from a cost function such as the translog. To examine the consistency of the data and determine the variation in cost shares among specialties, the cost ratios were calculated for a subset of all expenses for all 15 specialties separately for solo and group practitioners. The group cost shares are listed in Table I (A & B) and the solo cost shares in Table II (A & B).

Each ratio in these tables is the indicated expense item over the sum of all indicated expenses: these expenses include office expenses, materials, equipment, and labor which is divided into five categories, clerical and office staff, registered nurses, licensed practical nurses, technicians and physician extenders. Below each expense item is the number of the question on the 1976 physician's practice cost survey which is the source of the data. The expense total excludes malpractice insurance, auto expenses, physician

employee wages and, most importantly, physicians' net income. These variables were excluded due to the preliminary nature of this initial inquiry. These data are meant only to be suggestive. Nevertheless, it appears that variation in cost shares does exist among specialties.

Even more important, this initial data inquiry reveals some puzzling potential internal contradictions in the survey. First, total non-physician wages (question 52) is not equal to the sum of the wage payments to the individual labor categories (multiplied by 52 weeks). Second, for solo practitioners, the sum of all expenses including malpractice and automobile costs (which could be calculated easily because solos have only one member per firm) does not match total deductions.

Also, listed at the bottom of the tables are the average number of office visits per hour of patient contact and the average total number of hours worked per week.

Finally, Table III compares cost shares for both group and solo for general surgery and internal medicine, which are two of the most heavily represented specialties; these cost shares include expenses for auto use and malpractice insurance but not physician net income. A Chi-square test indicated that a significant difference exists at the 99% confidence level in the distribution of cost shares between general surgery and internal medicine for solo (72.4 16.8) and for group (77.25 16.8). This difference seems substantially attributable to malpractice insurance which is, as expected, much higher for surgeons.

TABLE I A

COST SHARES AND RELATED VARIABLES
FOR PARTNERSHIPS, UNINCORPORATED AND INCORPORATED GROUPS

SPECIALTY	1	2	3	4	5	6	7	8
OFFICE SPACE (Q. 42)	0.150	0.260	0.210	0.277	0.188	0.228	0.196	0.265
EQUIPMENT (Q. 42)	0.049	0.064	0.024	0.100	0.044	0.031	0.033	0.183
MATERIALS (Q. 42)	0.115	0.074	0.094	0.119	0.156	0.125	0.136	0.063
CLERK (Q. 53)	0.264	0.309	0.499	0.310	0.308	0.351	0.325	0.388
REG. NURSE (Q. 53)	0.234	0.084	0.096	0.054	0.119	0.118	0.097	0.046
L.P. NURSE (Q. 53)	0.035	0.037	0.043	0.020	0.072	0.048	0.055	0.008
TECHNICIANS (Q. 53)	0.138	0.144	0.033	0.114	0.085	0.069	0.136	0.027
EXTENDERS (Q. 53)	0.014	0.027	0.011	0.006	0.025	0.031	0.021	0.020
SUM OF Q. 53	0.651	0.601	0.672	0.504	0.612	0.616	0.633	0.489
COST SUM (Q. 42+53)	$149053	126231	134225	114242	127622	113200	147086	121901
DEDUCTIONS (Q. 61)	$344379	377189	374852	253232	218309	255763	302361	311477
WAGE DIFF. (Q. 53-52)	$ 8969	6989	15303	- 2209	6792	5428	13374	- 627
VISITS/HOUR (C.29A/27A)	3.66	2.65	5.84	2.41	4.46	4.75	2.91	3.01
HOURS/WEEK (Q. 28)	50.23	68.46	49.46	65.50	65.62	65.28	64.39	65.89
N =	35	37	28	59	117	137	108	63

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TABLE I B

COST SHARES AND RELATED VARIABLES
FOR PARTNERSHIPS, UNINCORPORATED AND INCORPORATED GROUPS

SPECIALTY	11	12	13	14	15	16	17
OFFICE SPACE (Q. 42)	0.218	0.176	0.216	0.261	0.223	0.391	0.307
EQUIPMENT (Q. 42)	0.042	0.088	0.050	0.058	0.022	0.042	0.049
MATERIALS (Q. 42)	0.117	0.127	0.141	0.066	0.158	0.021	0.092
CLERK (Q. 53)	0.341	0.398	0.409	0.343	0.288	0.424	0.320
REG. NURSE (Q. 53)	0.136	0.040	0.047	0.107	0.129	0.021	0.105
L.P. NURSE (Q. 53)	0.074	0.031	0.009	0.018	0.077	0.011	0.067
TECHNICIANS (Q. 53)	0.060	0.074	0.095	0.043	0.076	0.041	0.046
EXTENDERS (Q. 53)	0.013	0.067	0.033	0.104	0.028	0.049	0.014
SUM OF Q. 53	0.623	0.609	0.593	0.614	0.598	0.545	0.552
COST SUM (Q. 42+53)	$113084	120396	143642	120384	110866	89109	95362
DEDUCTIONS (Q. 61)	$265725	220072	359362	214783	191343	241775	225579
WAGE DIFF. (Q. 53-52)	$ 10238	24507	8069	2887	7840	-10653	2590
VISITS/HOUR (Q.29A/27A)	4.62	3.22	4.15	5.03	4.05	1.44	4.12
HOURS/WEEK (Q. 28)	65.31	52.90	65.95	54.21	61.48	59.10	62.79
N =	216	42	79	42	165	38	72

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TABLE II A

COST SHARES AND RELATED VARIABLES
FOR SOLO PRACTITIONERS AND SOLO EXPENSE SHARERS

SPECIALTY	1	2	3	4	5	6	7	8
OFFICE SPACE (Q. 42)	0.220	0.274	0.261	0.281	0.196	0.259	0.236	0.274
EQUIPMENT (Q. 42)	0.033	0.064	0.046	0.086	0.046	0.032	0.042	0.122
MATERIALS (Q. 42)	0.201	0.089	0.130	0.120	0.194	0.117	0.147	0.035
CLERK (Q. 53)	0.267	0.349	0.377	0.380	0.320	0.387	0.350	0.484
REG. NURSE (Q. 53)	0.132	0.076	0.054	0.050	0.092	0.118	0.086	0.069
L.P. NURSE (Q. 53)	0.046	0.036	0.049	0.014	0.063	0.039	0.041	0.003
TECHNICIANS (Q. 53)	0.070	0.095	0.054	0.070	0.060	0.039	0.074	0.013
EXTENDERS (Q. 53)	0.030	0.017	0.027	0.000	0.029	0.009	0.024	0.000
SUM OF Q. 53	0.546	0.573	0.562	0.513	0.565	0.592	0.575	0.569
COST SUM (Q. 42+53)	$55678	37992	38270	33333	38130	33196	36098	37074
DEDUCTIONS (Q. 57)	$56062	41826	46219	32855	37091	34263	33638	44404
DEDUCTION * DIFFERENCE	$-4483	-1870	2951	-5456	-6578	-8678	-6905	-7950
WAGE DIFF. (Q. 53-52)	$ 1054	1600	2313	-7856	255	1766	2146	2727
VISITS/HOUR (Q. 29A/27A)	4.57	2.43	4.66	2.72	4.30	3.83	2.97	2.60
HOURS/WEEK (Q. 28)	51.71	66.60	46.93	58.65	60.67	64.09	63.55	63.10
N =	109	62	97	34	289	247	205	70

*DEDUCTION DIFFERENCE = DEDUCTIONS (Q.57) - SUM (Q.42 + Q. 53 + Q.44D1 + Q. 45)

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TABLE II B

COST SHARES AND RELATED VARIABLES
FOR SOLO PRACTITIONERS AND SOLO EXPENSE SHARERS

SPECIALTY	11	12	13	14	15	16	17
OFFICE SPACE (Q. 42)	0.286	0.258	0.231	0.233	0.265	0.473	0.369
EQUIPMENT (Q. 42)	0.039	0.137	0.107	0.164	0.018	0.022	0.061
MATERIALS (Q. 42)	0.094	0.116	0.142	0.094	0.181	0.038	0.101
CLERK (Q. 53)	0.332	0.356	0.358	0.345	0.276	0.375	0.286
REG. NURSE (Q. 53)	0.148	0.038	0.027	0.048	0.145	0.017	0.082
L.P. NURSE (Q. 53)	0.048	0.020	0.016	0.042	0.053	0.003	0.042
TECHNICIANS (Q. 53)	0.029	0.033	0.101	0.047	0.039	0.008	0.029
EXTENDERS (Q. 53)	0.025	0.045	0.018	0.026	0.023	0.063	0.031
SUM OF Q. 53	0.581	0.490	0.520	0.509	0.536	0.467	0.469
COST SUM (Q. 42+53)	$32985	39293	49795	41631	31873	12371	36832
DEDUCTIONS (Q. 57)	$41289	49058	47854	37872	33583	16542	41843
DEDUCTION* DIFFERENCE	$-2907	3864	-13944	-12724	-2532	1305	-4022
WAGE DIFF. (Q. 53-52)	$ 1471	139	2703	-461	1436	-4624	1182
VISITS/HOUR (Q. 29A/27A)	3.70	3.56	3.48	3.90	3.94	1.21	3.49
HOURS/WEEK (Q. 28)	60.44	51.90	58.75	53.12	58.81	51.78	57.96
N =	139	109	55	103	175	374	57

*DEDUCTION DIFFERENCE = DEDUCTIONS (Q. 57) - SUM (Q. 42 - Q. 53 + Q. 44D1 + Q. 45)

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TABLE III

COST SHARES
FOR GENERAL SURGERY AND INTERNAL MEDICINE

	SOLO		GROUP	
SPECIALTY	6	7	6	7
OFFICE SPACE (Q. 42)	0.201	0.210	0.171	0.175
EQUIPMENT (Q. 42)	0.025	0.037	0.023	0.029
MATERIALS (Q. 42)	0.090	0.131	0.093	0.120
MALPRACTICE (Q. 42)	0.182	0.061	0.203	0.073
AUTO (Q. 45)	0.045	0.049	0.049	0.045
CLERK (Q. 53)	0.299	0.311	0.263	0.286
REG. NURSE (Q. 53)	0.091	0.077	0.088	0.085
L.P.NURSE (Q. 53)	0.030	0.037	0.036	0.048
TECHNICIANS (Q. 53)	0.030	0.066	0.052	0.120
EXTENDERS (Q. 53)	0.007	0.021	0.023	0.019
SUM OF Q. 53	0.457	0.512	0.451	0.559

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WORKPLAN AND METHODOLOGY

TASK I - DEVELOPMENT OF AN EQUITABLE MEDICARE ECONOMIC INDEX: AN ANALYSIS OF VARIATIONS IN PRACTICE COSTS AMONG SPECIALTIES

Contract 500-78-0052

TK No. 78-3162

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WORK PLAN FOR TASK I ANALYSIS

I. Overview

During the late sixties and early seventies, physician fees for services to Medicare patients increased at a sufficiently rapid pace to warrant the attention of both Congress and the Administration. In 1972, this attention produced an amendment to the Social Security Act, designed to control the rate of growth in these expenditures for physician services. The proposed method of controlling fees focused on restricting the rate of increase of the prevailing charge rates, which are the maximum allowable reimbursement rates allowed physicians for Medicare services. Up to that time, increases in actual charges had determined the growth in the prevailing charge rates, but with this amendment, the growth in these rates was to be determined independently by a newly specified Medicare-Economic Index and not by the growth in actual charges.

The prevailing charge rates are a part of the "usual, customary and reasonable" (UCR) charge rate system which sets reimbursement rates for Medicare physician services. Under this system, the physician receives the lowest of the actual, customary or prevailing charges given certain reasonable allowances for variation in the service provided. The customary charge (Level I) is the physician's usual charge or more specifically the median charge for that service during the preceding screen year. Customary charges lag current charges by at most eighteen months due to the lag in

sampling for the screen year.

Prior to the introduction of the Medicare Economic Index, Medicare carriers calculated the prevailing charge (Level II) for a given service by setting it equal to the 75th percentile of all customary charges, each charge weighted by its frequency. The carrier grouped customary charges for similar specialists within a given locality. Under this system, increases in the actual charge increased the customary charge, albeit with a lag, and also increased the prevailing charge if most physician charges were growing as indeed has been the case. Thus, the system had a built in feedback link for inflation in actual charges. Prevailing charges, which had been designed as the usual ceiling for reimbursement, were reflecting steady increases in practice charges. The Medicare Economic Index was introduced in 1976 to bound these increases in prevailing charges by specifying an independent growth rate for charges based on cost and thereby breaking this connection between the changes in actual and prevailing charges.

In its form, the Medicare Economic Index resembles a forward Laspeyres price index in which the ratio of total price change (current to base period) equals the sum of the changes in the input price ratios (also current to base period), each ratio weighted by its own share of base period expenditure. If the input price ratios measure the change in the opportunity cost of factor inputs to the physician firm, then their properly weighted sum, which is the Index itself, represents the percentage change

in the cost of physician output. Each base-prevailing-charge-rate is then multiplied by the Index. If only one Index is calculated, each prevailing charge increases by the same percentage and the relative position of each charge remains the same. Thus far, due to the absence of detailed data, only a single national Index has been computed annually for updating the charges of all specialties in all regions. Also due to the absence of data, BLS subindices acted as proxies for the actual input prices with cost share weights coming from a separate survey. In this formulation and particularly with these data, the Index has decisively broken the feedback link from actual charges to prevailing charges.

The following discussion will review the organization for Task I analysis, whose work product will assist in the further development of the Index. The discussion outlines the proposed work in each of these areas:

1. the specification of the form of the Index;
2. the estimation of the weights for the Index;
3. the exploration of the relationships among cost components;
4. the determination of separate specialty and regional groupings;

During the first project year, preliminary work will also be undertaken in the following areas:

- the selection of prices or price proxies, and the comparison of BLS indices and comparable survey data;
- the examination of the impact of the Index on the physician's choice of inputs to the firm and services to Medicare patients.

II. Detailed Work Plan

1. Price Indices

An index number measures the change in the magnitude of a variable from one situation to the next. These situations usually consist of different dates but can consist of different places or different groups. The variable is usually but not always a composite of other variables. The most commonly referenced type of index is the price index and the most widely cited of the price indices is the BLS consumer price which is a forward Laspeyres index. As indicated above, the Medicare Economic Index has the form of a Laspeyres price index.

The forward Laspeyres price index is the ratio of the sum of current prices (p_{1i}), each weighted by base quantities (q_{0i}) to the sum of base prices (p_{0i}), each weighted by base quantities (q_{0i}). Stated otherwise, the Laspeyres index equals the sum of the ratio of current (p_{1i}) to base prices (p_{0i}), each ratio weighted by its own share of base period expenditure:

$$\frac{\Sigma\, p_{1i}\, q_{0i}}{\Sigma\, p_{0i}\, q_{0i}} = \Sigma\, \frac{p_{1i}}{p_{0i}}\, \frac{p_{0i}\, q_{0i}}{\Sigma p_{0i}\, q_{0i}}$$ LASPEYRES

A closely related alternative to the forward Laspeyres is the forward Paasche index in which the base quantity weights are replaced by current quantity weights:

$$\frac{\Sigma\, p_{1i}\, q_{1i}}{\Sigma\, p_{0i}\, q_{1i}} = 1 \Big/ \left(\Sigma\, \frac{p_{1i}}{p_{0i}}\, \frac{p_{1i}\, q_{1i}}{\Sigma p_{1i}\, q_{1i}} \right)$$ PAASCHE

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The Paasche index requires the collection of data for current period weights in each successive period; on the other hand, the Laspeyres index only requires base period weights. Consequently, constructing a Paasche index entails considerably more effort and the resulting survey costs usually limit its use. However, cost considerations address only one criterion by which to evaluate the relative merit of different index forms.

The Laspeyres and Paascne indices are only two of the infinitely large class of index number forms. To select a price index from members of this class requires a set of criteria to evaluate both their measurement accuracy and their correspondence to economic assumptions. Irving Fisher (1922) devised two categories of three tests each to assess the measurement accuracy of index numbers. These tests will be reviewed briefly and applied to the Laspeyres index, the Paasche index, and Fisher's ideal index, which is the geometric mean of the first two. The Fisher index has the form

$$\left(\left(\frac{\Sigma P_{1i} Q_{0i}}{\Sigma P_{0i} Q_{0i}} \right) \left(\frac{\Sigma P_{1i} Q_{1i}}{\Sigma P_{0i} Q_{1i}} \right) \right)^{\frac{1}{2}}$$

Fisher's first category included three simple tests of measurement consistency. The first of these is the identity test which requires that if a situation (e.g. a year) is compared to itself, it equals one. The second test is the units test which states that a change in the measurement unit of money will not change the value of the index itself. The third is the homogeneity or proportionality test which requires that if all prices change by an equal proportion from base to current period, the index itself will change by that

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proportion. (This test is quite useful for it is consistent with the property of demand functions that they are homogenous of degree zero in both prices and income.) The Laspeyres, Paasche, and Fisher-ideal index numbers all pass each of these three tests. The demonstrations are straightforward.

The second category of Fisher's tests determines whether an index number is transitive. The first of these tests requires that the forward price (quantity) index equals the inverse of the backward price (quantity) index, or in other words, that the product of the index measured from current to base and of the index measured from base to current equals one. This test insures that a change in the reference point of measurement does not change the order of the index number values. Satisfaction of this consistency test would be particularly important if the index were comparing groups; if the test were not satisfied, switching from one reference group to another might not preserve the ordering of the index number values. Among the index forms being considered, only the Fisher ideal index satisfies this test (the component subscripts having been dropped for clarity):

$$\left(\frac{\Sigma p_1 q_0}{\Sigma p_0 q_0} \cdot \frac{\Sigma p_1 q_1}{\Sigma p_0 q_1} \right)^{\frac{1}{2}} \left(\frac{\Sigma p_0 q_1}{\Sigma p_1 q_1} \cdot \frac{\Sigma p_0 q_0}{\Sigma p_1 q_0} \right)^{\frac{1}{2}} = 1$$

However, neither the Laspeyres nor the Paasche index numbers satisfy this test; in the forward Laspeyres case:

$$\left(\frac{\Sigma p_1 q_0}{\Sigma p_0 q_0} \right) \left(\frac{\Sigma p_0 q_1}{\Sigma p_1 q_1} \right) \gtrless 1$$

(i.e. not necessarily equal) because the inverse of the forward Laspeyres is

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the backward not the forward Paasche, the latter appearing as the second term above.

The second test in this category requires a transitivity across components. In the case of a price index measured across time, this test specifies that the index measured from the base to the current period equal the product of the same index measured from the base to an intermediate period and from the intermediate period to the current period. None of the indices being considered satisfies this test; in the Laspeyres case:

$$\left(\frac{\Sigma p_t \; q_0}{\Sigma p_0 \; q_0}\right) > \left(\frac{\Sigma p_k \; q_0}{\Sigma p_0 \; q_0}\right)\left(\frac{\Sigma p_t \; q_k}{\Sigma p_k \; q_k}\right)$$

(The inequality takes account of substitution due to price changes; the right hand side allows some substitution and the left none.) Of Fisher's six tests this is the most difficult one to satisfy. More significantly, the test is somewhat unrealistic in that it supposes an index could be developed which would use only current and base prices to measure the full price change over the entire period. An example might explain this objection to the test: if the values in a price vector shift from the base to the intermediate period (k) and then shift back to the base vector values in the current period (t), then the left hand side above (whether Laspeyres, Paasche or Fisher-ideal) would equal unity-would show no change-but would not necessarily equal the right hand side; the right hand side would reflect the effect of the sequence of interim price changes.

This test has a simple but important consequence for inter-group comparisons.

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In comparing group A to group C, one should not derive that comparison by the indirect route of comparing A to B and then B to C if the index number chosen fails this test. The test has another more important consequence and that is pointing to the desirability of taking into account the interim sequence of price changes from the base to the current period (e.g., calculating the right not the left hand side above). Doing this instantaneously would produce the only well-known index which satisfies this test: the continuous Divisia index. Its discrete usable counter-part is the chained Laspeyres index which has been shown above not to satisfy this transitivity test.

The final Fisher test requires that the product of the price and quantity indices equals the change in total expenditure. Neither the Laspeyres nor the Paasche indices satisfy this test: in the case of the Laspeyres:

$$\left(\frac{\Sigma p_1 q_0}{\Sigma p_0 q_0}\right)\left(\frac{\Sigma p_0 q_1}{\Sigma p_0 q_0}\right) \gtrless \frac{\Sigma p_1 q_1}{\Sigma p_0 q_0} \quad .$$

However, the Fisher ideal index does

$$\left(\frac{\Sigma p_1 q_0}{\Sigma p_0 q_0}\,\frac{\Sigma p_1 q_1}{\Sigma p_0 q_1}\right)^{\frac{1}{2}}\left(\frac{\Sigma p_0 q_1}{\Sigma p_0 q_0}\,\frac{\Sigma p_1 q_1}{\Sigma p_1 q_0}\right)^{\frac{1}{2}} = \frac{\Sigma p_1 q_1}{\Sigma p_0 q_0}$$

This test is often referred to as the factor reversal test; note that the test requires that the total expenditure change is explained as the product of the price and quantity indices. This consistent separation of price and quantity changes can have useful applications if one is interested in isolating the price change component.

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Other indices satisfy these tests. Stuvel (1957) has suggested a formulation using only the forward Laspeyres price and quantity indices. His derivation started with a separation of the change in total value ($\Sigma p_{i1}q_{i1} - \Sigma p_{i0}q_{i0}$) into price and quantity changes; in this derivation, Stuvel averaged prices in the quantity difference equation and averaged quantities in the price equation. This is the essence of the difference between the Fisher ideal index and the Stuvel index although he did not highlight it. Whereas Stuvel weigns each price difference ($p_1 - p_0$) by its own average quantity ($q_1 + q_0$)/2), Fisher's index weighs each price ratio (p_1/p_0) by the product of its base (current) expenditure share and the total Paasche (Laspeyres) price change. Thus, the Fisher ideal index weighs the individual price ratios by the product of own expenditure shares and the change in overall prices. The Stuvel index separates price and quantity changes explicitly (denoted I_p and I_Q respectively below):

$$I_Q = \frac{L_Q - L_P}{2} + \left(\left(\frac{L_Q - L_P}{2}\right)^2 + \frac{\Sigma p_{i1} q_{i1}}{\Sigma p_{i0} \Sigma q_{i0}}\right)^{\frac{1}{2}}$$

$$I_p = \frac{L_p - L_Q}{2} + \left(\left(\frac{L_p - L_Q}{2}\right)^2 + \frac{\Sigma p_{i1} q_{i1}}{\Sigma p_{i0} \Sigma q_{i0}}\right)^{\frac{1}{2}}$$

where L_Q is the Laspeyres quantity index and L_p the price index. I_p is the total effect of all price changes and I_Q is the total effect of all quantity changes. Stuvel showed that his indices satisfied all the tests which Fisher's ideal index satisfied; in particular note that $I_Q I_p = (\Sigma p_1 q_1)/(\Sigma p_0 q_0)$.

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The above analysis applies several measurement tests to selected price indices. The remainder of this Section evaluates the economic implications of some indices. The basic Task I objective is the development of a price index for physician firms. If a cost function exists for physician firms, then the price index should reflect the structure of the cost function. Unfortunately, only a small portion of the literature on price indices analyzes the correspondence between cost functions and price indices. Two exceptions are Vartia (1976) and Diewert (1976, 1978).

A basic observation about the Laspeyres index is in order before considering the work of Vartia and Diewert. In the Laspeyres index, the total price change equals the ratio of the sum of the products of base quantities and current prices to the sum of the products of base quantities and base prices. The use of base quantities as weights in the current period implies that no substitution occurs in production as a result of the price changes. This implication points to a very restrictive assumption about the nature of the underlying production function; namely, that at least in the short-run (between the base and current periods), the elasticity of substitution between inputs is zero because the production function is characterized by fixed factor coefficients.

On the other hand, comparing current prices and quantities to base quantities and current prices, the Paasche index asks the hypothetical questions: what would have been the values of production at base prices given the current mix of quantities? The Paasche index assumed that complete substitution among quantities has occurred. Consequently, the Fisher index, which is the product of the Laspeyres and Paasche, brackets the substitution possibilities. If no substitution occurs, then the Fisher index reduces to the Laspeyres and with complete substitution to the Paasche.

Diewert developed the link between cost functions and the Vartia and Törnquist indices. Diewert's development is rather technical and only some results will be described. He defined a general criterion for the equivalence of the price index and the ratio of the cost function evaluated at current prices to the cost function evaluated at base prices. This criterion stated that the cost function is exact for a price index if given the base and current price vectors, the dual linearly homogeneous aggregator (i.e., 'production') function has a maximum which corresponds to the input base and current quantity vectors respectively.

Diewert then demonstrated that the Vartia index is exact for the Cobb-Douglas unit cost function. The Vartia index is defined as

$$\ln P_v = \Sigma \frac{(p_{1i}q_{1i} - p_{0i}q_{0i}) / (\ln(p_{1i}q_{1i}) - \ln(p_{0i}q_{0i}))}{(\Sigma p_{1i}q_{1i} - \Sigma p_{0i}q_{0i}) / (\ln(\Sigma p_{1i}q_{1i}) - \ln(\Sigma p_{0i}q_{0i}))} \ln \frac{p_{1i}}{p_{0i}}$$

and the Cobb Douglas unit cost function is defined as

$$\ln c = \alpha_0 + \Sigma \alpha_i \ln p_k$$

where c is cost per unit, α_0 a constant and p_i the input prices and $\ln P_v = \ln c_1 / \ln c_2$. The Vartia quantity index only substitutes the ratio of quantities in the cost term. The product of the Vartia price and quantity indices equals the change in total cost; in other words, the Vartia indices satisfy the Fisher factor reversal test. Furthermore, if the firm's technology corresponds to a linearly homogenous Cobb-Douglas production function (i.e., the elasticity of substitution equal to one), then the use of the Vartia index insures that the price ratios are weighted by factors consistent with the underlying tech-

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nology. With Cobb-Douglas technology, use of the Laspeyres index would over-report the price change while use of the Paasche index would under-report it.

Diewert goes on to show that the Törnquist index is exact for the translog unit cost function. This result is quite interesting because the translog function is much more flexible than the Cobb-Douglas, allowing non-constant elasticities of substitution. (As Theil (1973) noted the index is not unique to Törnquist and is frequently used as the discrete approximation to the Divisia index.) The Törnquist price index is

$$\Sigma \tfrac{1}{2}\left((p_{i0}\, q_{i0}/\Sigma\, p_{i0}q_{i0}) + (p_{i1}\, q_{i1}/\Sigma\, p_{i1}\, q_{i1}) \right) \ln\, (p_{i1}/p_{10}).$$

The first two ratios are the expenditure shares in the base and current period respectively. The translog unit cost function corresponding to a linearly homogeneous translog production function is

$$\ln c = b_0 + \Sigma\, b_i \ln p_i + \tfrac{1}{2} \Sigma\Sigma b_{ij} \ln p_i\, p_j$$

where c is cost per unit and p_i the input prices. The price index equals the ratio of the unit cost function evaluated at current prices to it evaluated at base prices or if this price index is designated as $\ln P_T$, $\ln P_T = \ln C_1/\ln C_0$, where C_1 is evaluated at current prices and C_0 as base prices. The quantity index substitutes quantities for prices in the price ratio terms above but the expentiture shares are identical. The quantity index equals the ratio of the aggregator function evaluated at current quantities to the aggregator function evaluated at base prices.

The product of the price and quantity indices does not quite equal the change in total cost but as Theil (1973) showed, approximate it to the third order of smallness. This means that the Törnquist index does not quite satisfy the Fisher factor-reversal test. If the firm's technology corresponds to a linearly homogenous translog production function, then use of this index will insure that price ratios are weighted by factors consistant with the underlying technology. The translog function allows a very flexible specification for the technology, and includes the Cobb-Douglas as a special case (σ=1).

Diewert demonstrated the exact correspondence between the Törnquist index and the translog function under constant returns to scale. Diewert (1976) also showed that when the translog does not exhibit constant returns, the Törnquist index does not correspond uniquely to the translog but to a wider class of functions including the translog. In this general form of the translog, output enters not as y_0 in the base and y_1 in the current period, but as $(y_0 y_1)^{\frac{1}{2}}$ in both. (Diewert demonstrated this result for the more general non-homothetic case in which the "cross-effect" of input price with output enters the cost function.)

Note that the index has two other rather minor limitations. First, if a quantity is present in some periods but not others, the log operator in the quantity index will be undefined in years in which this quantity is absent. This problem does not exist in the price index. Secondly, unlike the Laspeyres, Paasche, and Vartia, this index is not quite consistent in aggregation: combining subindicies will not exactly yield the overall index.

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The index has a major non-technical limitation. It requires the estimation of base and current expenditure shares because the weights are averages of the two periods. Continual sampling to determine current expenditure shares entails considerable re-survey expense; as indicated above, the Paasche index encounters the same objection. Of course, use of the information from the last and current period avoids the "drifting" problem which occurs when only a Laspeyres index is used; the standard resolution for the problem is the chaining (i.e., multiplying) of up-dated Laspeyres indices together. Chaining the Törnquist indices would improve this solution.

Diewert's work suggests a way to avoid some of the costs of resurvey. If the underlying technology remains constant then the estimated coefficients of the translog unit cost function can predict the Törnquist index. The predicted Törnquist index would equal the ratio of the product of the estimated coefficients $\hat{\beta}$ and the current price vector P_1 over the products of the coefficients and the base price vector P_0. For this reason, we will estimate cost function parameters; the discussion concerning that estimation is in the next section. Without the same re-sampling which this procedure seeks to avoid, one could not be confident that the underlying technology had not changed and that the estimated coefficients used in the ratio's numerator were correct. However, one might be somewhat confident that a substantial shift would not occur within, for example, five years which is the usual period for changing the base in the use of the Laspeyres indices. In fact, use of the annual re-survey data from the physician's practice cost survey to estimate cost functions for sequential years may suggest whether significant changes in the elasticities of substitution will have occurred in the near term. Review of the literature may also assist this evaluation.

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However, if the possibilities for substitution among inputs in physician firms is not considerable, then the correspondence between the flexible translog and the Törnquist index may not be that important. Through our own estimation and particularly that of others we will attempt to determine some range on the elasticities of substitution. Note that if such substitution was very low or very slow, then use of a Laspeyres type index would not mis-respresent critical cost shares. As part of this work, we will attempt to compare the implicit substitution possibilities represented by the Törnquist index with that represented by the Fisher ideal and the Stuvel indices by using some of Uzawa's (1962) results representing the Allen partial elasticities of substitution in terms of cost function derivatives.

Up to this point, the discussion has reviewed both measurement and economic criteria by which to evaluate various index forms. These two criteria address the issue of consistency in indices. Furthermore, the economic discussion focused on the extent to which the price index weights accurately represented the factor substitution possibilities. This discussion reaches the issue of efficiency but not that of equity and the Task I subject is the formulation of an equitable Medicare Economic Index. To reach this issue requires some consideration of the meaning of equity.

In a practical context, which avoids abstract arguments about welfare functions, equity might merely mean the absence of discrimination amoung groups differing significantly in technology. In this case, the groups might well be the separate specialties which would be treated alike only if their production

technology, as measured by their estimated elasticities of substitution, were not significantly different. If for some a priori reason particular groups were known to be inefficient in the use of inputs so that the estimated coefficients did not reflect an efficient technology, then consideration could be given to the use of other proxies for the true parameters in order ultimately to encourage efficiency. The application of this equity rule is then consistent with the efficient use of inputs in production. To an economist, this application would seem desirable because it represents the movement toward or preservation of Pareto optimality (recognizing of course the considerable market imperfections in the area of physician firms which point to second-best solutions).

On the other hand, establishing equity might mean not only avoiding discrimination among groups but also limiting the growth rate in Medicare fees to the growth rate of prices in another sector. This seems to be the equity goal because of the method chosen to define the fee limits through the Index. The current formulation of the Medicare Economic Index uses the growth rate in input prices from other sectors as a limit for the implied growth rate in our sector input prices. Through the Index, this limit on input prices translates into a limit on fees. How effectively the chosen means will reach the implicit goal is not clear and we will examine this issue in general terms. This examination will consider possible goals behind the index: for example, the maintenance of parity in the growth rate between total Medicare physician expenditures and earned income. Finally, the Goal of the index may not be explicit parity between growth rates but may merely be the establishment of some limits by which to control growth in fees and this possibility will be included in the discussion.

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2. Estimation of the Weights

The previous Section reviewed the structure and the merits of several indices. Every index requires a set of prices and estimated cost shares, and this Section addressed the problem of the estimation of the cost shares.

With the Laspeyres price index, only an estimate of the base expenditure or cost shares is necessary because the Laspeyres assumes the unimportance of the substitution between the base and current periods. On the other hand, the Paasche index needs the estimate of current cost shares. The Fisher Ideal, Stuvel (implicitly), Vartia and Törnquist indices all require estimates of cost shares in current and base periods. As indicated at the end of the last Section, estimation of the Cobb-Douglas and translog unit cost functions provide indirect measures of the Vartia and Törnquist indices respectively. The discussion in this Section includes a description of the proposed cost function estimation.

One can approach the estimation of cost shares directly by calculating the share of total expenditures in each of the several categories of input cost. The Physician Practice Cost Surveys in 1975 and 1976 include physician firms' annual expenses for all major input categories: office and auto expenses, medical equipment and materials, malpractice insurance, non-physician labor and physician net income. (The 1975 Survey contains much more detailed administrative expenses; and the Abt report has extensively analyzed these expenses.)

For all specialties, the tables in Appendix I present the cost shares of these categories excluding physician net income. Each ratio entry in the tables is

the dollar amount expended in the indicated category over the total dollar amount expended in all categories, excluding non-physician income. Consequently, each ratio is the sum of costs in the indicated category to the sum of all non-physician costs where each sum is taken over all physicians in a given specialty. These ratios weight the cost distributions of the larger firms more heavily and therefore have bias positively related to the firm's total cost. A more appropriate ratio estimator is the average of the ratios, rathern than the ratio of the averages, because the former does not weigh firms differently. If within a given specialty, the population (not the sample) cost share parameters differs due to size of output, different prices, etc., then only regression analysis can capture these differences and that approach is introduced below. Note that these ratios presented in the Appendix tables and the chi-square statistic were intended to be only convenient and descriptive of the differences among specialties.

The appropriate test to determine simultaneous equality of the cost share patterns across specialties is a multivariate analysis of variance. (Anderson, 1958) This test consists of multiple analyses of variance, each analysis of variance testing the equality of a given mean (e.g., a particular type of cost share - labor, materials) across all specialties and the multiple analysis testing the equality of all these sets of means simultaneously. (Because the n cost shares sum to one, testing the equality of the n-1 shares is sufficient.) Therefore, using a multiple analysis of variance would test whether significant differences existed in the distribution of all cost ratios across all specialties. Performing the test directly on the ratios is desirable because the ratios enter directly into the price index formulation. However, ratios can be biased estimators of the relation between the two variables if a regression line between the two variables does

not pass through the origin but has a constant term. This possibility might well occur in the relation between a cost component and the total cost; when the total is positive, some components can indeed be zero (if those inputs are unused in production).

The problems with ratio estimators leads to a consideration of alternatives. The next most naive approach is to average the cost components per physician across all firms in a given specialty and test whether the distribution of these average costs differed across specialties, again using a multivariate analysis of variance. If opportunities for economies of scale were constant across firms and related to the number of physicians per firm yet this number on average differed across specialties, then the specialities with larger firms would show smaller cost component averages, ceteris paribus, but the ratios would be identical across firm size because the distribution of cost across categories has by assumption not changed. However, even this test on the component averages would not locate the cause of inter-firm or inter-specialty differences; in the example given, all-around lower costs could have resulted from lower prices. Establishing a causal relationship between changes in total costs (or its components) and other factors requires introducing a model to explain variation in form costs.

An appropriate model is the cost function of the firm which relates total cost to output and the set of input prices. The cost function contains the same information about production as its dual the transformation function which describes the efficient production possibility set, if the cost function is non-decreasing and strictly positive when output is nonzero and if the cost function is concave and linearly homogeneous in it input prices for any given level of output.

Several problems arise in the estimation of cost functions for physicians. If the firm does not cost minimize, then the estimated parameters of any given cost function will not represent the possibilities for substitution under conditions of efficient production. This outcome makes interpretation of the estimated parameters difficult if not impossible. Note that if the producer in a non-competitive industry is a cost-minimizer, identifying the parameters of the firm in the non-competitive case may still be a problem because equilibrium price and quantity will be a function of both supply and demand. Another way to view this problem is that the system is not identified by estimation of the cost function alone. A further problem is the endogeneity of the price of the physician's own labor; to interpret the estimated parameters as determining elasticities of substitution requires that the prices be exogenous to the firm. Similarly, output should be exogenous. In the case of physician firms, neither output nor price of physician labor is likely to be exogenous; in fact, they are probably jointly determined.

As a mathematical function, a cost function can assume a variety of forms. Two commonly used forms are the Cobb-Douglas and the translog. Each corresponds to different assumptions about the possibilites for substitution among factor inputs although their differences extend beyond this because the translog allows non-homotheticity and a greater flexibility in the form of non-constant returns to scale. The Cobb-Douglas form implies that the Allen partial elasticities of substitution between pairs of inputs ($_{ij}$) is constant and equal to one, while the translog form imposes no restrictions: the elasticities can assume any

real value and can vary with the level of the cost shares and the relationship between a cost share and other input prices. The greater flexibility of the translog usually recommends its use although its estimation is more costly. The following discussion will review both the forms of the Cobb-Douglas and the translog function, discuss some of the estimation problems and conclude with a few comments on the use of the cost functions in the price index.

The Cobb-Douglas production function has the form:

$$\ln Y = k + \Sigma \alpha_i \ln X_i$$

where Y is output, and the X_i the set of inputs. Cost minimization subject to this production function (and through a series of substitutions) yields a total cost function, whose estimated form is

$$\ln (C-P_j) = \alpha + \Sigma \alpha_i \alpha_Y \ln(P_i-P_j) + \alpha_Y \ln Y$$

where C is total cost, each P_i is the price of the input, each α_i is from the production or function and α_Y is the inverse of the sum of the α_i. If the sum of the α_i is less (greater) than one, decreasing (increasing) returns to scale are present. The product $\alpha_i\alpha_Y$ equals the cost share of the ith input (p_iX_i/C). In order that the estimated parameters of all cost shares sum to one $\Sigma p_iX_i/C = 1$, the estimation of the cost function must incorporate the equivalent restriction $\alpha_i\alpha_Y = 1$ which is implemented by subtraction of any one price from total cost and all other prices.

Information on one or more prices may not be available; for example, the Survey does not have price data on office materials, equipment or floorspace or medical materials or equipment. If these prices were correctly assumed to be constant across firms, then their effect would enter the estimated constant term but since several prices are missing, their estimated cost shares would only be known in the aggregate as $1 - \Sigma \alpha_i \alpha_\gamma$ where $\Sigma \alpha_i \alpha_\gamma$ covers only those inputs with non-missing price data. In fact, the only prices available are the wage rates for labor. Thus, even if the unavailable prices were constant across firms, no unique cost shares could be inferred. (The Cobb-Douglas function has a redeeming feature here; its assumption of strong separability in inputs (all $\sigma_{ij} = 1$) is equivalent to the existence of a consistent aggregate subindex of prices so that if one price could be found which would describe the entire set of missing prices no information would be lost, but unforunately no such price index is in the Survey or conveniently recomends itself.) Because the missing prices are non-labor, one might be tempted to treat them all as "capital" and use the price of external financing as a proxy but this use confuses the real and the financial sides of the firm's operation; for example, if two firms faced the same price vector of inputs but different prices for external financing because of imperfect capital markets, then their choice of input might well differ but not because of different possibilities for substitution in production.

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The translog cost function has the form

$$\ln C = a_0 + a_Y + \tfrac{1}{2} g_{YY} (\ln Y)^2 + \sum_i a_i \ln P_i +$$

$$\tfrac{1}{2} \sum_i \sum_j g_{ij} \ln P_i P_j + \sum_i \sum_j g_{Yi} \ln Y \ln P_i$$

where C is total cost, Y output and P_i the input prices. For the translog function to have the desirable characteristic of homogeneity of degree one in prices, which implies that a uniform increase in all prices results in a proportionate increase in total costs, the following restrictions must hold:

$$\sum_i a_i = 1 \; ; \; \sum_i g_{Yi} = 0; \; \sum_i g_{ij} = \sum_j g_{ij} = \sum_i \sum_j g_{ij} = 0$$

These restrictions constrain the cost share sum to equal one (Christiansen and Greene, 1976).

Differentiating the cost function with respect to prices yields a set of cost share equations of the form

$$S_i = a_i + g_{Yi} Y + \sum_j g_{ij} \ln P_j$$

where the S_i are the cost shares of the i th input. Because of the probable substantial multicollinearity among all the price and cross-price variables in the cost function, estimation of the cost share equations yields considerably more precise parameters. However, the cost share equations do not include several parameters appearing in the cost function itself. Their absence prevents testing hypotheses of homotheticity and homogeneity which respectively imply $g_{Yi} = 0$ and $g_{Yi} = g_{YY} = 0$. If these hold and if all the $g_{ij} = 0$, then the translog reduces to the Cobb-Douglas form. Joint estimation of the cost function and the cost share equations allows the testing of these hypothe-

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ses. Because the cost shares must sum to one, one cost share equation must be dropped from this system. To derive parameter estimates which are invariant to the equation being dropped requires maximum-likelihood estimation; iterating the Zellner procedure (IZEF) is one approach (Christensen and Greene, 1976).

The cost function considered above included only one output. Multiple outputs may more accurately describe the physician firm. Specifying multiple outputs for the translog function requires adapting the log-operator of output because of the possibility that one or more of the outputs for some firms might be zero. Caves, Christensen and Trathaway (1978) have proposed use of the Box-Cox metric $f(y) = (y^k - 1)/k$ to replace the log metric. The log metric remains for prices and total cost allowing easier parameter restrictions to test homogeneity of degree one in prices. Their generalized translog multiproduct cost function (GTMCF) has the form

$$\ln C = a_0 + \sum_i a_i (Y_i^k - 1)/k + \sum_i b_i \ln P_i + \tfrac{1}{2} \sum_i \sum_j g_{ij} \ln P_i \ln P_j +$$

$$\tfrac{1}{2} \sum_i \sum_j d_{ij} (Y_i^k - 1)/k \, (Y_j^k - 1)/k + \sum_i \sum_j r_{ij} (Y^k - 1)/k \ln P_j$$

where the Y_j are outputs, P_j prices and C total cost. Non-linear estimation is required to determine which value of k minimizes the sum of squared errors.

Extensive data unavailability affects both the translog and the Cobb-Douglas but may limit the use of the computationally more expensive translog (and particularly the GTMCF) in the estimation of the cost function. Only prices for non-physician labor appear in the Survey. Estimation of a cost function with only the available prices will create serious econometric problems with omitted

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variables; only if the omitted variables are uncorrelated with the included variables will the parameter estimates of the included variables not be biased. Assuming this lack of correlation, the estimated elasticities of substitution among labor inputs would be unbiased.

The estimate of the intercept will still be biased even if the omitted and included variables are uncorrelated, a particular problem in the translog cost share equation where the intercept identifies the impact of own price on the share; intercept bias will not disappear unless the omitted variable were irrelevant which by assumption it isn't (Kmenta, 1971). Only if the omitted prices were constant across all firms, which is extremely unlikely will their absence make no difference at all.

Some of the variables included in the Survey present problems. The first problem occurs with the specification of output. We will review the literature to assess which unitary measures of output have been used. The actual output is a potential change in the patient's health status, which is unmeasureable; the most commonly used output measure is simply the number of visits. Some investigators have used output measures such as total revenue which would be inappropriate in the cost function. In the estimation of the cost function, we will attempt to correct for some of the endogeneity of the output measure through two-stage least squares.

Disaggregating output could point to the construction of a case-mix index which recognized that each physician produces a different mix of diagnoses and treat-

ments. The supplementary work in Task 1 on Case-Mix Analysis indicates the direction of our exploratory efforts. This Analysis points to difficulties in linking the NAMCS data to the Survey data (especially via Survey question 35). Using a unitary measure of output which obscures the true variation in the multiple outputs may bias the estimated coefficients on the price of physician own labor because that price is probably positively correlated with case-mix components requiring more effort (e.g., more difficult diagnoses and treatments). The Case-Mix Analysis does not address one basic characteristic of both physician input and output; namely, its stochastic nature. Physicians can't accurately control the mix of problems presenting themselves or the outcomes from the treatments they recommend. The variation across specialties in malpractice expenses indicates among other things that uncertainty in treatment outcomes varies substantially. How this uncertainty affects the physician's production function is not clear.

Disaggregating output presents an interesting possibility for assessing the impact of the Index. The case-mix approach characterizes output as the mix of diagnoses and treatments. Another approach would be to characterize output as the mix of patient types, specifically Medicare and non-Medicare. If the Index creates an effective ceiling, then the coefficient on the cross-product term could indicate the extent to which substitution away from Medicare patients occurred.

The other major problem is the measurement of the price of physician labor. The price of physician labor is not exogenous to the firm; the physician who organizes a practice will decide simultaneously an expected output and an

al unit but measuring it is difficult. A proxy might be the fee for the most flexible of the services offered, the return visit, adjusted by its average length and reduced to exclude administrative expenses. In the long run, the private opportunity cost might be defined differently: the potential physician can choose among occupations and the practicing physician can choose between practice settings (e.g., a salaried government or HMO position would be alternatives).

Consideration of the correct price of physician time raises a related issue. That issue is the proper model for the physician firm. Some investigators have modeled the firm as if it were only an individual physician supplying own labor and others as if it were a firm. If the firm is the appropriate model, then estimation of the "supply of labor" function is a mis-specification even in the case of the solo-practitioner. The correct specification would be the firm's derived demand for physician labor function which should include firm output, the price of own input and the price of all alternative inputs for all observations even though some of the firms do not use some of the inputs. Determining which is the appropriate model is difficult; use of specification tests in estimation might help.

Another serious problem in the Survey is the extent of measurement error in the variables. The tables in Appendix I suggest the extent of the problem. In a regression, errors in variables create bias in the estimated coefficients unless in the limit the value of the term on the left vanishes:

$$\text{plim}(1/n(X'(u-VB))) = \text{plim}(1/n\ X'u) - \text{plim}(1/n\ \tilde{X}'V)B - \text{plim}(1/n\ V'V)B$$

where n is the sample size, X the measured and $\tilde{X}$ the true matrix of the values

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of the variables, u the vector of the residuals in the regression without error and V, the vector of the measurement errors (Johnston, 1972). Even if the first term on the left vanishes because u is uncorrelated with the true value of X and the measurement error ($X = \tilde{X} + V$), the third term will not vanish because the measurement error itself will be positive in the limit. Also, the second term will possibly be positive in the limit because the measurement error will likely be correlated with the true values of X; for instance, the measurement error in physician gross and net income probably increases with the true value of these variables.

The estimation of the translog or Cobb-Douglas cost function would be very useful because of their direct role as analogues to the Törnquist or Vartia price indices. Unfortunately, estimation of the appropriately specified cost function seems unlikely given the available data. Some alternatives might be investigated. The cost shares might be related to available measures of output, input prices, and other exogenous variables characterizing the structure of the local market for health 'production'. Such a set of equations would not correspond to a cost function but would specify a set of 'behavioral' cost share functions, whose stability is unknown because of their essentially ad hoc character.

However, if variables are selected for their ability to explain variation in the cost shares, then those which enter will minimize the residual given the current characteristics of the physician, the practice and its locale. If the cost share estimates are derived as the product of the estimated coefficients and the characteristics of the individual practice, then these

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estimates will reflect the existing local market. Thus, the more ad hoc the specification, the greater the likelihood that the cost share estimates will allow the Index to maintain market characteristics, including inefficiencies. For instance, if more non-competitive areas (e.g., those with both greater than average fees and physician density) showed cost shares weighted more heavily toward more costly inputs (e.g., physician input), then the Index would permit these areas a greater increase in fees. To remedy this problem would require dropping the variable from the estimation or substituting another value for the independent variable for these firms.

To draw the threads of this discussion to a conclusion, we propose a three-tiered method of determining the cost share weights. First, we will calculate the overall cost shares directly as averages of the individual shares, testing to determine if significant differences exist among specialties (c.f. Section 3). Second, we will estimate the Cobb-Douglas cost function with available prices constraining the labor shares to equal one less the shares of the unincluded variables and acknowledging the serious specification errors which omission of the other variables would create. We will also estimate the translog for limited portions of the sample. Third, we will estimate ad hoc cost share equations which include variables other than price and output.

Collection of the missing price data in future years would allow estimation and use of the cost function in the Index. However, determining the price of many inputs is difficult. An alternative would be to sample not the prices but the quantity of inputs and estimate the production function instead of the cost function. The price index could then be approximated indirectly

by using the result that the product of the price and quantity indices equals the ratio of the current to base expenditure or explicitly by calculating the ratio of two other ratios: the ratio of current to base expenditures over the ratio of current to base quantity index. For those indices which don't satisfy the factor reversal test exaclty (e.g., the Törnquist), this calculation would only be approximate.

3. Cost Component Relationships

The previous Section presented methods to calculate the weights used in the Medicare Economic Index. This Section briefly reviews related work aimed at further examination of the nature of the cost components. This work includes the examination of the levels of the components, their relation to net physician income, the rates of change in those levels and the relationship between other factors and both the levels and the cost shares.

For each specialty grouping, we will calculate the mean for each cost component per physician. The cost components will include the same categories as listed in Section 2: office and auto expenses, medical equipment and materials, non-physician labor, malpractice insurance, and physician net income. (These are the listed categories on the 1975 and 1976 Physician Practice Cost Survey.) Dividing each component by the number of physicians in the practice before averaging across all physicians in the same specialty will partially normalize the cost components by producing values which are costs per physician.

This normalization does not remove the effect of possible economies of scale. If opportunities for economies of scale were constant across specialties and related to the number of physicians in the practice but at the same time the average number of physicians per practice differed across specialties, then the costs per physician would also differ across specialties. However, if the analysis in Section 2 indicates what other researcher have shown, namely, that at most only modest economies of scale are present and are exhausted with fairly small practices, then one might apply the same analysis suggested in

the beginning of Section 2 to determine whether significant differences among specialties exist.

To determine the equality of each set of cost components simultaneously, we will again use a multivariate analysis of variance test. This approach will test a joint hypothesis, a composite of several individual hypotheses. Each individual hypotheses specifies that the means for a given cost component are equal across all specialties. The test for that equality is a one-way analysis of variance. The test for the simultaneous equality of the means across specialties for the several cost components requires a multivariate analysis of variance. We will calculate the appropriate F-statistic to test whether significant differences exist (Anderson, 1958).

For all specialties grouped together, we will also perform a multivariate analysis of variance to determine if cost components vary across size of practice. We will measure size of practice by number of physicians and cost components by cost per physician. Again we will test a joint hypothesis, each individual hypothesis specifying that the means for a given cost component are equal across all practice sizes. This analysis will supplement the work in Section 2 and will help determine whether costs per physician decline in some intermediate range of practice size.

To examine the relationship of physician net income to total other costs, we could rely on a simple correlation between the two. We could posit a simple behavioral relation which would argue that if physician net income were under the firm's control, then it might be explained in a multiple regression by

output measures and total non-physician costs. We also can determine whether the relation between physician net income and total non-physician cost were correlated by using the Wilcoxon-rank test, the rank test being useful because of the possible bias in reporting errors for both costs. However, none of these ad hoc tests reaches the issue of whether high cost firms exist, firms whose costs are not justified by the level and mix of their output. A better approach would be to examine the pattern of the estimated residuals in the estimated total cost equation to determine if the size or variance of the estimated residuals increases for those firms having large total costs. If the translog functions were estimated, one could also examine the estimated residuals for these same firms to determine the pattern of estimated cost shares.

Both the 1975 and the 1976 Surveys contain questions relating to the physician firm's cost components. The question numbers and the cost categories to which they refer are the following:

	1975	1976
OFFICE EXPENSES	57-1,3,60,61	42-1
AUTO EXPENSES	57-2	45
MEDICAL EQUIPMENT	57-4	42-2
MEDICAL MATERIALS		42-3
MALPRACTICE INS.	62	44D
PHYSICIAN EMPLOYEES	68	48A
NONPHYSICIAN EMPLOYEES	68	53
TOTAL DEDUCTIONS	74	57,61
NET INCOME	62	62
GROSS INCOME	72,73	55,58

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To measure the change in cost components for each specialty, we will compute both the change in individual cost components (75 to 76) divided by base year cost component (75) and this same change divided by base year total cost. The sum of the second percentage changes will equal the overall change in costs for a specialty. We can again apply a multivariate analysis of variance to the second percentage change. (Because the variable is a ratio, some caution should be exercised in interpreting the results.) To determine whether changes in cost components have been similar across components for any given specialty, we can we can apply a one-way analysis of variance to the set of changes for each specialty. A more complete test for structural shift in cost shares from 1975 to 1976 would utilize the Chow-test applied to the pooled estimation and to the separate estimation of cost functions for each year. We will consider the time series problems involved in pooling.

Finally, we will examine the impact of other factors on the estimation of cost shares. To do this, we will include other variables in the estimation of the translog cost share equations discussed in Section 2. These variables would include measures of the physician's individual background (e.g., age), the physician's practice (e.g., mixture of patient types, diagnostic services requested), or the area's characteristics in which the practice is located (e.g., measures of physician density by type of physician, number of hospital beds). In Section 2, we indicated that both output and the physician wage are endogenous to the firm, causing biased estimates of their coefficients in the cost share regressions. A potential remedy replaces the variables with their predicted values using two-stage least squares where excluded exogenous variables and included exogenous variables are used in the first stage to estimate the

predicted values. The excluded exogenous variables will include some of the variable indicated above. Entering these variables through the two-stage least squares process does not require making a priori assumptions about the strong separability of these variables and the other included variables in the cost equation itself. If the strong separability assumption is made, some of these variables could be separately included in the cost equation itself.

4. Specialty and Regional Groupings

Currently the Medicare Economic Index does not distinguish among specialties or regions but consists instead of a single index with one set of cost shares. Using this single index and implicitly ignoring these differences among specialties creates potential problems because the Index may generate increases in prevailing fees which do not correspond to increases in cost, particularly for those specialties whose cost shares are far from the average.

However, if these inter- or even intra-specialty differences in cost shares do no reflect differences in technology, the estimation and use of separate cost share weights will reflect other influences. Estimating the parameters of the cost function from the survey data presents some considerable difficulties as indicated above in Section 2. The Survey data is deficient but even were the data complete and available, the resulting estimated parameters of the cost functions would represent the impact of market distortions. Work to date on testing the relative economic efficiency of firms has dealt with the case of allegedly undistorted markets in which all firms are price-takers (Yotopoulos and Lau).

In the apparent absence of a definable test to determine relative economic efficiency, we propose two ad hoc procedures. First, we will perform an analysis of covariance to test for significant differences among specialties. The analysis of covariance requires finding the estimated residual for the pooled sample (SSE_p) and for the classes within the sample ($\sum_{i=1}^{P} SSE_i$) and computing the following test distributed as an F-statistic with indicated degrees of freedom.

$$\frac{SSE_p - \sum_{i=1}^{p} SSE_i/k(p-1)}{\sum_{i=1}^{p} SSE_i/p(m-k)}$$

where k is the number of variables, p the number of classes, and $\hat{m}$ the number of observations per class. (The degrees of freedom specification is different if there are different numbers of observations per class.) This analysis of covariance will be applied directly to the estimated cost functions using the specialties as the classes. We will first test for overall homogeneity across all specialties. If that test indicates no significant differences across classes, then the testing would stop and it could be assumed that the single index is adequate. If significant differences are indicated, we will then divide the sample into surgical and non-surgical specialties, repeating the test for each of these two samples. If significant differences again exist, we will then divide the non-surgical sample into general/family practice and remaining specialties. The division into classes at each step should follow some notion of different practice organizations; for example, several surgical specialities may show no significant differences in cost shares and may be arguably treated as a single class. (Because of the sequential testing we will consider some adjustment to the degrees of freedom.)

Second, this test could be used within specialties if the specialties were partitioned into classes on the basis of cost per visit. If we accept the hypothesis that specialties with lower cost per visit constituted more "efficient" firms in some sense, then we could test for the differences in

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cost functions between the lower and higher cost firms. If no difference existed then the cost shares of the specialty (or group of specialties) could be used but if significant differences were indicated we could consider using the cost shares of the lower cost partition. These differences might be attributable to greater efficiencies in the lower cost firms or to other influences including market distortions.

Because of limited samples within several of the specialties, we will be unable to test for simultaneous homogeneity across region-specialty pairs unless we restrict those pairings to a reasonable minimum. A more inclusive set of pairings would be possible were we to pool the 1975 and 1976 Surveys which we will consider doing. On the other hand, we could allow the regional variations to enter only through price variation, which seems to be more plausable in the sense that within a given specialty or group of specialties the technology of physician practices should not vary substantially across regions. If we puruse this course, we may want to test to determine whether the sample shows significant differences among the regions in cost per visit functions, rather than total cost functions, by including regional dummy variables in this type of cost function. If significant differences appeared, that would warrant use of regional prices or use of adjusted national prices. One source for that adjustment might be the estimated coefficients themselves.

DATA ANALYSIS FOR PROJECT AREA I

Cost Shares Estimates

One objective of this project is the determination of the actual cost shares of each expense category facing the physician firm. These cost shares will function as the weights in the Medicare Economic Index.

The most fundamental approach to determine these cost shares is to calculate the ratio of each expense item to total expenses for each specialty. This method assumes that no significant variation in cost shares occur within the specialty and that significant variation does exist among specialties. To test the first assumption requires estimation of cost share equations as derived, for instance, from a cost function such as the translog. To examine the consistency of the data and determine the variation in cost shares among specialties, the cost ratios were calculated for a subset of all expenses for all 15 specialties separately for solo and group practitioners. The group cost shares are listed in Table I (A & B) and the solo cost shares in Table II (A & B).

Each ratio in these tables is the indicated expense item over the sum of all indicated expenses: these expenses include office expenses, materials, equipment, and labor which is divided into five categories, clerical and office staff, registered nurses, licensed practical nurses, technicians and physician extenders. Below each expense item is the number of the question on the 1976 physician's practice cost survey which is the source of the data. The expense total excludes malpractice insurance, auto expenses, physician

employee wages and, most importantly, physicians' net income. These variables were excluded due to the preliminary nature of this initial inquiry. These data are meant only to be suggestive. Nevertheless, it appears that variation in cost shares does exist among specialties.

Even more important, this initial data inquiry reveals some puzzling potential internal contradictions in the survey. First, total non-physician wages (question 52) is not equal to the sum of the wage payments to the individual labor categories (multiplied by 52 weeks). Second, for solo practitioners, the sum of all expenses including malpractice and automobile costs (which could be calculated easily because solos have only one member per firm) does not match total deductions.

Also, listed at the bottom of the tables are the average number of office visits per hour of patient contact and the average total number of hours worked per week.

Finally, Table III compares cost shares for both group and solo for general surgery and internal medicine, which are two of the most heavily represented specialties; these cost shares include expenses for auto use and malpractice insurance but not physician net income. A Chi-square test indicated that a significant difference exists at the 99% confidence level in the distribution of cost shares between general surgery and internal medicine for solo (72.4 16.8) and for group (77.25 16.8). This difference seems substantially attributable to malpractice insurance which is, as expected, much higher for surgeons.

TABLE I A

COST SHARES AND RELATED VARIABLES
FOR PARTNERSHIPS, UNINCORPORATED AND INCORPORATED GROUPS

SPECIALTY	1	2	3	4	5	6	7	8
OFFICE SPACE (Q. 42)	0.150	0.260	0.210	0.277	0.188	0.228	0.198	0.265
EQUIPMENT (Q. 42)	0.049	0.064	0.024	0.100	0.044	0.031	0.033	0.183
MATERIALS (Q. 42)	0.115	0.074	0.094	0.119	0.156	0.125	0.136	0.063
CLERK (Q. 53)	0.264	0.309	0.499	0.310	0.308	0.351	0.325	0.388
REG. NURSE (Q. 53)	0.234	0.084	0.096	0.054	0.119	0.118	0.097	0.046
L.P. NURSE (Q. 53)	0.035	0.037	0.043	0.020	0.072	0.048	0.055	0.008
TECHNICIANS (Q. 53)	0.138	0.144	0.033	0.114	0.088	0.069	0.136	0.027
EXTENDERS (Q. 53)	0.014	0.027	0.011	0.006	0.025	0.031	0.021	0.020
SUM OF Q. 53	0.681	0.601	0.672	0.504	0.612	0.616	0.633	0.489
COST SUM (Q. 42+53)	$149053	126231	134225	114242	127622	113200	147086	121901
DEDUCTIONS (Q. 61)	$344379	377189	374852	258232	218309	255763	302361	311477
WAGE DIFF. (Q. 53-52)	$ 8989	6989	15303	- 2209	6792	5428	13374	- 627
VISITS/HOUR (Q.29A/27A)	3.66	2.65	5.84	2.41	4.46	4.75	2.91	3.01
HOURS/WEEK (Q. 28)	50.23	68.46	49.46	65.50	65.62	65.28	64.39	65.89
N =	35	37	28	59	117	137	108	63

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TABLE I B

COST SHARES AND RELATED VARIABLES
FOR PARTNERSHIPS, UNINCORPORATED AND INCORPORATED GROUPS

SPECIALTY	11	12	13	14	15	16	17
OFFICE SPACE (Q. 42)	0.218	0.176	0.216	0.261	0.223	0.391	0.307
EQUIPMENT (Q. 42)	0.042	0.088	0.050	0.058	0.022	0.042	0.049
MATERIALS (Q. 42)	0.117	0.127	0.141	0.066	0.158	0.021	0.092
CLERK (Q. 53)	0.341	0.398	0.409	0.343	0.288	0.424	0.320
REG. NURSE (Q. 53)	0.136	0.040	0.047	0.107	0.129	0.021	0.105
L.P. NURSE (Q. 53)	0.074	0.031	0.009	0.018	0.077	0.011	0.067
TECHNICIANS (Q. 53)	0.060	0.074	0.095	0.043	0.076	0.041	0.046
EXTENDERS (Q. 53)	0.013	0.067	0.033	0.104	0.028	0.049	0.014
SUM OF Q. 53	0.623	0.609	0.593	0.614	0.598	0.545	0.552
COST SUM (Q. 42+53)	$113084	120396	145642	120384	110866	89109	95362
DEDUCTIONS (Q. 61)	$265725	220072	359362	214783	191343	241775	225579
WAGE DIFF. (Q. 53-52)	$ 10238	24507	8069	2887	7840	-10653	2590
VISITS/HOUR (Q.29A/27A)	4.62	3.22	4.15	5.03	4.05	1.44	4.12
HOURS/WEEK (Q. 28)	65.31	52.90	65.95	54.21	61.48	59.10	62.79
N =	216	42	79	42	165	38	72

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TABLE II A

COST SHARES AND RELATED VARIABLES
FOR SOLO PRACTITIONERS AND SOLO EXPENSE SHARERS

SPECIALTY	1	2	3	4	5	6	7	8
OFFICE SPACE (Q. 42)	0.220	0.274	0.261	0.281	0.196	0.259	0.236	0.274
EQUIPMENT (Q. 42)	0.033	0.064	0.046	0.086	0.046	0.032	0.042	0.122
MATERIALS (Q. 42)	0.201	0.089	0.130	0.120	0.194	0.117	0.147	0.035
CLERK (Q. 53)	0.267	0.349	0.377	0.380	0.320	0.387	0.350	0.484
REG. NURSE (Q. 53)	0.132	0.076	0.054	0.050	0.092	0.118	0.086	0.069
L.P. NURSE (Q. 53)	0.046	0.036	0.049	0.014	0.063	0.039	0.041	0.003
TECHNICIANS (Q. 53)	0.070	0.095	0.054	0.070	0.060	0.039	0.074	0.013
EXTENDERS (Q. 53)	0.030	0.017	0.027	0.000	0.029	0.009	0.024	0.000
SUM OF Q. 53	0.546	0.573	0.562	0.513	0.565	0.592	0.575	0.569
COST SUM (Q. 42+53)	$55678	37992	38270	33333	38130	33196	36098	37074
DEDUCTIONS (Q. 57)	$56062	41826	46219	32855	37091	34263	33638	44404
DEDUCTION * DIFFERENCE	$-4483	-1870	2951	-5456	-6578	-8678	-6905	-7950
WAGE DIFF. (Q. 53-52)	$ 1054	1600	2313	-7856	285	1766	2146	2727
VISITS/HOUR (Q. 29A/27A)	4.57	2.43	4.66	2.72	4.30	3.83	2.97	2.60
HOURS/WEEK (Q. 28)	51.71	66.60	46.93	58.65	60.67	64.09	63.55	63.10
N =	109	62	97	34	289	247	205	70

*DEDUCTION DIFFERENCE = DEDUCTIONS (Q.57) - SUM (Q.42 + Q. 53 + Q.44D1 + Q. 45)

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TABLE II B

COST SHARES AND RELATED VARIABLES
FOR SOLO PRACTITIONERS AND SOLO EXPENSE SHARERS

SPECIALTY	11	12	13	14	15	16	17
OFFICE SPACE (Q. 42)	0.286	0.258	0.231	0.233	0.265	0.473	0.369
EQUIPMENT (Q. 42)	0.039	0.137	0.107	0.164	0.018	0.022	0.061
MATERIALS (Q. 42)	0.094	0.116	0.142	0.094	0.181	0.038	0.101
CLERK (Q. 53)	0.332	0.356	0.358	0.345	0.276	0.375	0.286
REG. NURSE (Q. 53)	0.148	0.036	0.027	0.048	0.145	0.017	0.082
L.P. NURSE (Q. 53)	0.048	0.020	0.016	0.042	0.053	0.003	0.042
TECHNICIANS (Q. 53)	0.029	0.033	0.101	0.047	0.039	0.008	0.029
EXTENDERS (Q. 53)	0.025	0.045	0.018	0.026	0.023	0.063	0.031
SUM OF Q. 53	0.581	0.490	0.520	0.509	0.536	0.467	0.469
COST SUM (Q. 42+53)	$32985	39293	49795	41631	31873	12371	36832
DEDUCTIONS (Q. 57)	$41289	49058	47854	37872	33583	16542	41843
DEDUCTION* DIFFERENCE	$-2907	3884	-13944	-12724	-2532	1305	-4022
WAGE DIFF. (Q. 53-52)	$ 1471	139	2703	-461	1486	-4624	1182
VISITS/HOUR (Q. 29A/27A)	3.70	3.56	3.48	3.90	3.94	1.21	3.49
HOURS/WEEK (Q. 28)	60.44	51.90	58.75	53.12	58.61	51.78	57.96
N =	159	109	55	103	175	374	57

*DEDUCTION DIFFERENCE = DEDUCTIONS (Q. 57) - SUM (Q.42 - Q. 53 + Q. 44D1 + Q. 45)

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TABLE III

COST SHARES
FOR GENERAL SURGERY AND INTERNAL MEDICINE

	SOLO		GROUP	
SPECIALTY	6	7	6	7
OFFICE SPACE (Q. 42)	0.201	0.210	0.171	0.175
EQUIPMENT (Q. 42)	0.025	0.037	0.023	0.029
MATERIALS (Q. 42)	0.090	0.131	0.093	0.120
MALPRACTICE (Q. 44a1)	0.182	0.061	0.203	0.073
AUTO (Q. 45)	0.045	0.049	0.049	0.045
CLERK (Q. 53)	0.299	0.311	0.263	0.286
REG. NURSE (Q. 53)	0.091	0.077	0.088	0.085
L.P.NURSE (Q. 53)	0.030	0.037	0.036	0.048
TECHNICIANS (Q. 53)	0.030	0.066	0.052	0.120
EXTENDERS (Q. 53)	0.007	0.021	0.023	0.019
SUM OF Q. 53	0.457	0.512	0.451	0.559

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